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Structural wage rigidity in West Germany 1950 - 1989: Some new econometric evidence

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Some New Econometric Evidence

by

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Structural Wage Rigidity in West Germany 1950-1989. Some New Econometric Evidence

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1. Introduction

Is the wage structure too rigid? Since the mid-1970s, this question has been one of the most persistent themes of the economic policy debate in West Germany. In recent months, the question is asked with a new sense of urgency: after the political and economic unification of the country in 1990, a huge imbalance between the labour markets in the West and the post-communist East will be a fact of life for years to come so that common sense and economics speak in favour of a sustained interregional wage differentiation. However, actual wage agreements, which mostly envisage an equalization of wages between West and East by the mid-1990s, do not take this need for flexibility into account.

In this paper, we shall take a fresh look at the empirical evidence on structural wage rigidity for the last four decades of economic history in West Germany. We shall do so by applying some more recently developed techniques of co-integration, which provide a useful framework for the econometric analysis of structural labour market disequilibria and their wage effects. While the subject of our inquiry is the West German labour market since the early 1950s, the results do have implications for what one might expect for the future in a united Germany which has taken over basically all institutional features of the West's so-called social market economy, including the predominance of collective bargaining in wage determination.

The paper consists of six parts. After the introduction (Part 1), we shall briefly discuss some major theoretical issues which are linked to the definition of structural wage rigidity (Part 2). The basic framework for the subsequent empirical analysis is set in Part 3; it is then applied to interregional wage rigidity in Part 4 and to intersectoral/interindustrial wage rigidity in Part 5. The paper is concluded with a few remarks on the relevance of our analysis for the ongoing economic policy debate on wage differentiation in Germany (Part 6).

2. The Rigidity of the Wage Structure: Some Conceptual Issues

To make economic sense, the term 'wage rigidity' must be defined with reference to some sort of labour market disequilibrium. In the case of the wage level, the relevant disequilibrium is rather easy to identify as the economy-wide excess supply of labour and to approximate, e.g., by the overall unemployment rate. In the case of the wage structure, things are more complex. As there is a widespread confusion in the literature about the terminology and the concept of structural rigidities, it is worthwhile to pursue this matter in some detail.

Let us begin with four very general postulates on the wage structure of an economy:

(i) Whenever there is no Keynesian and no classical unemployment in the economy as a whole, there must be some equilibrium structure of wages which is compatible with full employment, i.e., which conforms to the multitude of full employment value productivities of labour all over the economy (neglecting cases of zero or negative labour productivity). Clearly, this is an implication of Walrasian general equilibrium theory under the usual set of convexity assumptions.

(ii) Technological progress and intersectoral shifts of demand in goods markets - not the least due to changes in the international division of labour - may lead to revaluations of human capital in the economy; thus the structure of value productivities of labour at full employment is likely to be changed and with it the equilibrium structure of wages. If the actual wage structure does not adequately reflect any such change, then occupational, sectoral and/or regional disequilibria may emerge. The extent and the speed of any subsequent adjustment of the actual towards the equilibrium wage structure can sensibly be called the degree of structural wage flexibility, or conversely, the degree of structural wage rigidity.

(iii) In a very general sense, the revaluation of human capital in the course of structural change in goods markets is a mere consequence of some sort of labour immobility: if workers were able to change occupations, sectors and/or regions <u>without</u> incurring frictional costs in the broad sense of the term, then no change of the wage structure would be required to preserve full employment. In practice, however, there will be costs - presumably moderate ones in the case of sectors, higher ones in the case of regions, and often prohibitive ones in the case of occupations. In any case, the degree of mobility is likely to be a function of time: the more time elapses after an exogenous productivity or terms-of-trade shock to the structure of human capital values, the more likely it is - ceteris paribus - that the full employment structure of wages returns to the standard of normality since the quantity movements of labour across sectoral, regional and/or occupational boundaries as a reaction to wage incentives gradually dissipate the rental element in the full employment wage structure. To the extent that labour mobility can be expected to remove the disequilibrium in the long run, the rigidity of the actual wage structure boils down to a short- and medium-term phenomenon. Hence, the relevant questions to ask are the following ones: is the wage structure flexible enough to accomodate shocks which lead to temporary structural disequilibria? Does it allow to save jobs in any sector hit hard by a negative shock and thus give the respective workforce the 'breathing space' to search for valuable alternative employment opportunities without becoming unemployed?

(iv) Of course, the alternative to labour moving to where the jobs are created is that the new jobs, i. e the new complementary capital stock, moves to labour. If we realistically assume a fairly perfect capital market, in which rates of return are more or less equalized across feasible uses, and if we further exclude the possibility of deliberate government manipulation of the price of capital through subsidies, then the only instrument to achieve a redirection of capital is the flexibility of the wage structure which - ceteris paribus - determines the relative profitability of investment between the relevant structural units. Clearly, the higher the costs of labour mobility, the more one has to rely on the movement of capital to reach again full employment all over the economy. In the extreme - say, e.g., with labour being completely immobile between regions - capital alone has to shoulder the burden of adjustment.

Prima facie, these four postulates look like fairly innocent truisms which could hardly arouse any controversy. After all, price and quantity movements are considered to be substitutes almost everywhere in economics. Nevertheless, the postulates have implications which are in conflict with some common arguments in the economic policy debate and - still more importantly - with the thrust of empirical tests of interindustrial and interregional rigidities of the wage structure.

As to economic policy, it is often claimed that structural wage flexibility of whatever kind[1] is likely to hinder or slow down structural change of employment: with wages reacting to sectoral disequilibria, the expanding sectors raise and the contracting sectors lower wages so that - compared to a situation with no wage differentiation - employment in the expanding sectors grows slower and employment in the contracting sectors shrinks slower as the 'wage burden' is distributed 'pro-structurally'.[2] This argument misses the message of our postulates (iii) and (iv) and thereby confuses cause and effect: only if the mobility of labour and capital is insufficient at the given wage structure to bring about the required intersectoral dislocation of labour will a change of this very structure be required as a substitute (and incentive) for higher mobility. Preventing this - purely instrumental - wage flexibility to emerge boils down to a strategy of pushing relatively immobile workers into a state of unemployment in which they are (assumed to be) ready to switch sectors. By implication, the workers regard the state of unemployment as less desirable than the state of employment at a lower wage in their 'old' sector since they leave the former, but do not leave the latter. The state of unemployment is thus used as a deliberate weapon to force

1 Usually, the case is made with respect to the intersectoral wage structure, but - with a few non-essential modifications - it carries over to other 'structural units', notably regions.

2 See, e.g., Bell, Freeman (1985), pp. 17-18, Franz (1989), pp. 309-311, and the Ifo-Institute, Munich, as quoted by Donges, Schmidt et al. (1988), p. 190.

reluctant workers into moving. In fact, this is the core of the famous 'Swedish-' or 'Rehn/Meidner-model' of structural change at a constant intersectoral wage structure[3] in which the state of unemployment is sweatened for workers by a partial socialisation of mobility costs.

In practice, which of the two strategies of structural change is preferred - the 'individualistic' one via wage differentiation or the 'socialist' one via government-sponsored unemployment - is largely a matter of ideological taste, not of economic substance: liberals with a preference for market solutions usually favour the former, social democrats with a preference for wage equalisation lean towards the latter. Note, however, that there is no apriori case for any of the two variants in terms of structural adjustment speed: it is a purely empirical question whether intersectoral wage flexibility or government-sponsored unemployment makes the labour force be relocated faster from shrinking to expanding sectors. To our knowledge, this question has not been tackled rigorously in the empirical literature on structural change. In any case, it is beside the point to argue that wage flexibility per se may hinder structural change of employment: clearly, the flexibility of the wage structure is a possible instrument of structural adjustment to be compared in its efficiency with other instruments such as a deliberate socialisation of mobility costs of the unemployed.[4]

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[3] For the Rehn-Meidner model, see the concise summary by Lundberg (1985), pp. 17-18.

[4] Some authors (notably Bell, Freeman 1985, pp. 17-18) argue that an explicit distinction should be made between two kinds of structural wage flexibility, namely the flexibility in the wake of an intersectoral shift of product demand and the flexibility in response to intersectoral differences in productivity growth which are due to purely technological reasons. The former is regarded to be conducive to overall employment growth in any case since it serves as an instrument to spur intersectoral movements of labour in the sense described above in the text; in contrast, the latter is considered to be ambiguous in its effect on employment growth since an above-average wage rise in the sector with fast productivity growth may dampen employment growth in this sector by more than the corresponding wage moderation in the sector with slow productivity growth helps to save jobs.
This analytical distinction between two qualities of wage differentiation ...

As to empirical testing, the question arises what kind of framework can and should be used to estimate the extent of structural wage rigidity in the sense defined above. Traditionally, 'wage flexibility' has been measured by simply calculating some proxy of the structural wage dispersion and then analyzing its changes over time. The results of this procedure have in general been quite unambiguous, at least for the case of West Germany:[5] all over the last twenty years, the wage dispersion among industries, between industry and services and between different levels of qualification of the workforce have been notably constant. If anything, a slight trend towards equalisation is discernible for the early 1970s and a slight reverse trend for recent years, but the thrust of the evidence points towards a constant structure. From this, it has been

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is misleading: in a standard neoclassical model with two sectors, which basically underlies all arguments on intersectoral wage flexibility and differentiation, <u>any</u> exogenous shock at given wages has an effect on sectoral labour demand <u>only via</u> the intersectoral ratio of marginal value productivities of labour, no matter whether the shock originates in technology and is transmitted through the ratio of physical productivities or whether it originates in goods markets and is transmitted through output prices. In either case, the sector experiencing the rise of value productivity has the scope to expand employment at the given wage and thus to attract labour by offering higher wages, while the sector with the relative loss comes under pressure to release workers; in either case, the adjustment requires a temporary wage differentiation to the extent that intersectoral labour mobility is less than perfect. Hence it makes no sense to link the employment effects of wage differentiation to the nature or type of the original exogenous shock.

Note also that, whatever the shock may have been in the first place, the <u>only</u> economic rationale for a subsequent wage differentiation is to preserve full employment in the case of insufficient intersectoral mobility at the given wage structure. Then, however, it is a moot question whether - without wage differentiation - more jobs could be created in the expanding sector than would be lost in the contracting one; after all, without appropriate wage incentives (or, for that matter, a Swedish-style subsidization of mobility), released workers in the contracting sector will not fill the job slots in the expanding part of the economy, and thus actual employment will be lower without than with intersectoral wage flexibility. Clearly, the employment performance of an economy should be measured by the standard of the actual number of jobs, not of the job potential, which would have been realised if workers had been more mobile.

5 See Breithaupt, Soltwedel (1980), Bell, Freeman (1985), Gundlach (1986), Soltwedel (1988), Soltwedel, Trapp (1988), Franz (1989), and - with a somewhat different focus of research - Fels, Gundlach (1990).

concluded that much of the lament about structural rigidities is grossly overdone.[6]

In our view, this conclusion is not warranted simply because the empirical tests do not address - let alone answer - the relevant question. What the evidence in fact shows is that there has been no sustained trend towards a flattening or a stratification of the wage structure. However, the evidence says virtually nothing about the crucial issue whether this very structure did or did not react to temporary or permanent disequilibria in specific labour (sub-)markets.

On a technical level, the testing design of this type of evidence is based on the implicit assumption that a rise (decline) of wage differentiation across sectors/regions in a descriptive sense can serve as a reliable indicator for wage flexibility in the economic sense defined above. Quite obviously, however, it cannot, as the example of structural change in the West German economy in the last 20 years easily shows: if, e.g., the real wage across industries at time t_0 is positively correlated with the physical labour productivity (which, in turn, is determined i.a. by the capital intensity of production!), and if the industries with relatively high physical labour productivity suffer intersectoral terms-of-trade losses in the course of structural change from t_0 to t_1, then the full employment wage structure with an immobile labour force becomes <u>less</u> dispersed, i.e., the high (physical) productivity industry should lose at least some of its wage lead. Note that this is basically what happened after 1973 with the partial decline of the West German heavy industries which are traditionally high productivity and high wage sectors. The same argument holds mutatis mutandis for regions if structural change favours a traditionally low wage area. Again, examples from the economic history of West Germany are easy to find: the decline of highly industrialized, high-wage Northrhine-Westfalia and the concommittant rise of (relatively) low-wage Bavaria may have called for <u>less</u> interregional wage differentiation to mitigate 'perverse' migration incentives.

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[6] See in particular Franz (1989), pp. 325-326.

Behind this 'technical' deficiency, there is a deep conceptual fallacy: as the usual measures describe the extent of wage differentiation without any consideration of sectoral/regional labour market disequilibria, they must by construction remain blind to the core of the matter, namely the link between sectoral/regional wage dynamics and sectoral regional labour market disequilibria. In short: they miss what wage flexibility really ought to mean.

3. A Framework for Empirical Analysis

A stylized model of intersectoral wage flexibility may clarify this crucial point and redirect the analysis into the right channels. Note that, in the following theoretical paragraph, we use the term 'sector' in a broad sense, meaning either sectors of economic activity ('industries') or regions. Let the process of wage determination in an economy be described by the simple equation

$$(1) \quad \log (w/p)_t = -\eta \, z_t + \epsilon_t$$

where w denotes the aggregate wage level, p the aggregate price level, z a measure of aggregate excess supply in the labour market[7] (possibly the unemployment rate or the log thereof), η a constant parameter ($\eta \geq 0$), ϵ a random error term, and the subscript t the point in time, i.e. with annual data the relevant year. Equation (1) may be regarded as a long-run wage equation in a bargaining-type model.[8] Let the corresponding wage equation for any sector i (i = 1, 2, ..., n) be given by

$$(2) \quad \log (w/p)_t^i = -\eta^i z_t^i - \theta^i (z^i - z)_t + \epsilon_t^i$$

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[7] For the moment, we simply assume that there is something like a relevant labour market, both for the economy as a whole and for any sector. We shall return to this important matter below.

[8] See, e.g., most papers in Calmfors (1990).

with the superscript i identifying the sectoral variables $((w/p)^i, z^i)$ and parameters $(\eta^i, \theta^i \geq 0)$. Let the sector i - whether an industry or a region - be small enough to make w/p and z independent of any single $(w/p)^i$ and z^i (for $i = 1, 2, \ldots, n$).[9] Equation (2) allows for different combinations of sectoral wage flexibility with respect to aggregate and sectoral labour market disequilibria. E.g., for $\eta^i > 0$ and $\theta^i = 0$, the sectoral real wage reacts exclusively to aggregate disequilibria, not to specifically sectoral ones, i.e. not to that part of z^i which goes beyond or falls short of z (meaning z^i-z); conversely, for $\eta^i = 0$ and $\theta^i > 0$, it is only the specifically sectoral disequilibrium that matters, and not the aggregate one; for $\eta^i = \theta^i$, the sectoral real wage reacts to just the 'whole' sectoral disequilbrium (z^i) which, however, may to some extent merely reflect an aggregate disequilibrium (z) as typically happens in cyclical upswings and downturns that hit the economy as a whole. Of course, any other combination of η^i and θ^i is feasible provided that η^i, $\theta^i \geq 0$. Intuitively, it is quite obvious that the key to the matter of intersectoral wage rigidity or flexibility lies in the parameter θ^i: if - in an appropriately designed econometric setting which specifies long-run as well as short- and medium-run dynamic effects -, the θ^i turns out to be zero or close to zero, the sectoral wage can be regarded as 'rigid' with respect to specifically sectoral labour market disequilibria. If θ^i turns out significant and relatively high, then a fair amount of structural wage flexibility can be diagnosed.

Although extremely stylized and simplified, the above two equations do catch the essence of what structural wage flexibility and rigidity should sensibly mean according to our line of thought in the preceeding paragraph. This can best be seen by combining (1) and (2) in the equation

$$(3) \quad \log\,(w/p)_t^i - \log\,(w/p)_t = -(\eta^i-\eta)z_t - \theta^i(z^i-z)_t + v_t$$

with $v_t = \epsilon_t - \epsilon_t^i$. Plausibly enough, equation (3) states that the real wage differential between sector i and the aggregate depends first on the

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9 This is not a crucial assumption. Dropping it would complicate the analysis, but not alter the main line of reasoning.

aggregate labour market disequilibrium z_t - to the extent that the sectoral and the aggregate real wage differ in flexibility with respect to z_t - and second on the specifically sectoral disequilibrium $(z^i-z)_t$ - to the extent that the sectoral real wage reacts to the sectoral disequilibrium.

Under two supplementary assumptions, which are fairly innocent, equation (3) can be further simplified: (i) let $p_t^i = p_t$, i.e. the price level with which the nominal wage is deflated, is assumed to be the same all across the country. For all that matters in practice, this restriction makes good sense: if p is taken to be the consumer price index, there are simply no separate indices defined and available for the various relevant sectors, be they industries or regions of an economy. As to the short-run dynamics of nominal wages in the wake of price changes which is not specified in equation (3), it is also quite reasonable to assume that there are no differences in the sectoral lag structures: after all, why should, say, Bavarian workers, firms, and unions be more or less inflation-sensitive than their counterparts in Hesse or Lower Saxony? Hence the intersectoral difference in real wages and their changes over time can quite safely be approximated by the intersectoral differential in nominal wage terms. (ii) In a similar vein, we can assume that $\eta^i = \eta$, i.e. the extent of wage flexibility/rigidity with respect to aggregate disequilibria is imposed to be the same all across the economy. Again, the rationale is intuitively plausible: why should we expect any relevant agent in, say again, Bavaria to react differently to a common aggregate shock than his/her counterpart in other regions? Apart from this a priori consideration, more practical econometric arguments call for imposing the restriction $\eta^i = \eta$: in most estimates of various specifications of this stylized model, it proved difficult enough to identify the parameter θ^i even with the restriction applied; in fact, a simultaneous identification of $(\eta^i-\eta)$ and θ^i turned out virtually impossible due to the all too limited informational content of the time-series data used. As far as the general form of (3) was taken as the basis for estimation, the results pointed towards an insignificant parameter $(\eta^i-\eta)$.

Under these two simplifying assumptions, equation (3) can be written as

$$(4) \quad \log w_t^i - \log w_t = -\theta^i(z^i-z)_t + v_t$$

which simply says that the deviation of the wage in sector i from the aggregate wage is a negative function of the difference in the degree of excess labour supply between sector i and the economy as a whole. It is important to recognize how the parameter θ^i should be interpreted, independent of the precise definitions of the variables w and z which will be discussed below. Taken at face value, θ^i measures the marginal effect of a change in the deviation of the sectoral extent from the aggregate extent of labour market disequilibrium on the (log-)deviation of the sectoral from the aggregate wage level. Whether this marginal effect is of a short- or medium-run or of a long-run nature crucially depends on the time-series properties of the two relevant variables $\log(w^i/w)_t$ and $(z^i-z)_t$. If these two variables are stationary, i.e. if they have a constant mean over time, θ^i boils down to a measure of short- or medium-run flexibility, independent of how the dynamics of the equation is specified for actual estimation purposes: after any exogenous shock changing $(z^i-z)_t$, the economy will eventually return to the long-run equilibrium wage ratio $(w^i/w)_t$. If the relevant variables are non-stationary, things are more complex: provided the variables are integrated of the same degree, say degree one (difference stationarity), short- <u>and</u> long-run effects may be discernible if the variables are co-integrated, i.e. if they share a stochastic trend. An estimation in levels such as in equation (4) would then become a test for co-integration, θ^i would have to be interpreted as a long-run coefficient; an additional equation which appropriately specifies the short- and the medium-run dynamics would then serve as a framework to estimate the short- and the medium-run effects of $(z^i-z)_t$ on $(w^i/w)_t$. We shall return to this matter below when discussing actual specification issues.[10]

Although a most useful analytical starting-point, equation (4) is unduely restrictive in one important way which deserves some closer examination. Rewriting (4), we obtain

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[10] For summaries of the economic meaning of the statistical concepts referred to in the text, see Granger (1986), Hendry (1986), and Stock, Watson (1988). Note that, of course, the variants described do not exhaust the range of possible time series constellations.

$$(5) \quad \log w_t^i = \log w_t - \theta^i(z^i - z)_t + v_t.$$

This reformulation makes clear that, if the variable $\log(w^i/w)_t$ in equation (4) is to be stationary and if both $\log(w_t^i)$ and $\log(w_t)$ are non-stationary, which is most likely to be the case in view of the rapid wage growth over any sample period of post-war economic history in West Germany, then $\log(w_t^i)$ and $\log(w_t)$ have to be co-integrated with a long-run coefficient which exactly equals one. Conversely, if $\log(w^i/w)_t$ turns out to be non-stationary, this may simply be a reflection of a false restriction on the long-run coefficient, which in reality is not equal to one. There are important economic issues hidden behind this formal matter: clearly, any changes of variables which determine wages in the long run and which are not included in the equation, may lead to a systematic deviation of the sectoral wage from the aggregate wage level, independent of any sectoral disequilibria. E.g., the average wage in region i may fall back relative to the economy-wide average, if - ceteris paribus - the industrial structure of the regional economy gradually shifts towards relatively low productivity branches; or the average wage in, say, industry i may increase if - again ceteris paribus - the average level of qualification and pay of the workforce in this industry rises relative to the economy-wide average. As far as the use of wage indices which correct for this kind of structural effects, is not feasible due to data limitations, some systematic long-run changes of the wage differential between an industry or region and the economy as a whole is very likely to be oberservable in the long-run.

To account for these effects, a generalized version of equation (4) and (5) is used, namely

$$(6) \quad \log w_t^i = \theta_\beta^i \log w_t - \theta_\gamma^i(z^i - z)_t + v_t$$

with θ_β^i ($\theta_\beta^i \geq 0$) and θ_γ^i ($\theta_\gamma^i \geq 0$) denoting the long-run co-integration parameters of $\log(w_t^i)$ with $\log(w_t)$ and $(z^i-z)_t$ respectively. Again, different cases may be distinguished: (i) $\log(w_t^i)$ and $\log(w_t)$ are non-stationary, but co-integrated. In this case, θ_β^i is the long-run co-integration parameter, so that $[\log(w_t^i) - \theta_\beta^i \log(w_t)]$ is stationary. If $(z^i-z)_t$ is also stationary, θ_γ^i measures the short- and medium-run effect of the specifically sectoral labour market disequilbrium on the (stationary) variable $[\log(w_t^i) - \theta_\beta^i \log(w_t)]$. If, in turn, $(z^i-z)_t$ is non-stationary, θ_γ will be zero; this follows from the fact that $\log(w_t^i)$ and $\log(w_t)$ are co-integrated so that $(z^i-z)_t$ - while itself integrated - cannot be co-integrated with either $\log(w_t^i)$ or $\log(w_t)$ or a linear combination of the two. (ii) All variables are non-stationary, and $\log(w_t^i)$ is co-integrated with the linear combination $[\theta_\beta^i \log(w_t) - \theta_\gamma^i (z^i-z)_t]$ where both θ_β^i and θ_γ^i measure long-run effects with a seperate equation defining the short-run dynamics. Other constellations may be thought of, but need not concern us here.

So much for the general framework of empirical analysis. Before we turn to the details of the model to be estimated, we have yet to clear up what the sectoral labour market disequilibrium should really mean and thus how (z^i-z) should in principle be proxied. To tackle this question, one has to make an explicit distinction between the two types of structural units considered in this paper, notably regions on the one side and sectors in the narrow sense (or 'industries') on the other.

As to regions, the answer to the question is straightforward and hardly controversial: as long as labour mobility between the relevant regions can be assumed to be low enough to justify the assumption of a more or less exogenous regional labour supply at any point in time, the region's unemployment rate (or the log thereof) can be taken as an adequate disequilibrium variable.[11] As a consequence, the specifically regional component of unemployment can be taken to be the difference (or, for that matter, the log-difference) of the particular regional and the national unemployment rate.

As to sectors and industries, the question is somewhat more complex. Clearly, a sector of economic activity like an industry does per se have no genuine labour market as have geographic entities like a country or a state. Instead, structural change evokes a persistent reshuffling of labour between industries, with the labour mobility at a given wage structure across the relevant units being much greater than in the case

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11 All throughout this paper, the unemployment rate is defined as the share of registered unemployed persons in the total labour force.

of regions. This is the corollary of saying that the degree of human capital specificity is much higher in regional than in sectoral terms. Hence, on a more practical empirical level, no such thing as a sector's labour supply can be reasonably defined and thus no unemployment rate as a proxy for the underutilization of that supply can be calculated. Nevertheless, industrial crises (booms) involving large-scale lay-offs (hirings) are plain facts of experience; and it is no moot question to ask whether and to what extent the short- or medium-term fluctuations of sectoral economic activity are accompanied by likewise wage movements which may temporarily mitigate the quantity adjustments.

It is important to realize that industrial or sectoral crises can hardly be anything else than short- or medium-term phenomena. In the long-run, a large-scale lay-off of industrial workers will lead either to their re-absorption in other sectors of the economy, or to a greater pool of long-term unemployed which then cannot be assigned any more to a specific industry. This is the very consequence of the economic logic of structural change: as soon as an industry or a sector has shrunk to a competitive size, its 'historical' workforce cannot be taken any more as a sensible reference standard defining an 'industrial' or 'sectoral' unemployment which, in the extreme, would be perpetuated indefinitely. Also, any rationale for an intersectoral wage differentiation to accomodate the process of structural change would disappear in the long-run. In statistical terms, all this means that any measure of industrial or sectoral labour market disequilibrium (z^i-z) must be defined relative to some 'normal' or trend path of structural change, i.e. (z^i-z) must by definition be trend stationary and its effect on the sectoral wage must be constrained to be zero in the long-run. We shall develop such-like measures below when specifying the actual model to be estimated.

In this context, it is essential to have a very precise idea of the subtle difference between intersectoral and interregional wage flexibility/ rigidity. After an industry has gone through a crisis and finally reached a new (smaller) competitive size, any economic rationale for this industry to pay lower wages than others to facilitate adjustment is also gone. Nevertheless, if the respective industrial crisis has been geographically concentrated and if, as a long-term consequence of this very crisis, a particular region is left with a permanently higher unemployment rate relative to some aggregate standard, then a long-term regional wage differentiation is called for. To pick a historical example: if the adjustment crisis in the steel industry all over Germany in the early 1980's had left the state of Northrhine-Westfalia, which hosts the Ruhr valley steel industry, with a permanently higher regional unemployment rate relative to the national average, an economic case can be made for a general long-term wage differentiation between Northrhine-Westfalia and the rest of the country, but not for a long-term wage differentiation between the steel industry per se and other industries. In a sense, Northrhine-Westfalia would have to 'devalue' its labour force across the board to attract the capital necessary to cut back its unemployment rate to the level before the crisis or to induce a corresponding emigration of a part of its workforce to other regions of the country.

4. Interregional Wage Rigidity

We now elaborate a more specific model for estimation purposes. To start with, we shall focus on interregional wage rigidity, i.e. we shall try to give an econometric answer to the question whether and to what extent the wage structure across regional units (in West Germany until 1989: the eleven states) has reacted in the past to specifically regional labour market disequilibria, be they temporary or permanent.

4.1. Model Specification

To arrive at a model which basically displays the desired long-run characteristics of (6), we assume that the wage at time t in any region i is determined within the general dynamic framework[12]

$$(7)\quad \log w_t^i = \alpha_0 + \alpha_1 \log w_{t-1}^i + \beta_0 \log w_t + \beta_1 \log w_{t-1} + \dots$$
$$\dots + \gamma_0 (z^i - z)_t + \gamma_1 (z^i - z)_{t-1} + \epsilon_t$$

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[12] See Hendry, Pagan, Sargan, (1984), pp. 1041-1049, and Rüdel (1989), pp. 28-37 with a discussion of various types of models which can be derived by imposing different sets of restrictions on the parameters of an equation of type (7).

with w^i defined as the wage in region i, w as the wage in the economy as a whole, and (z^i-z) as the specifically regional labour market disequilibrium. The superscript i denotes the region, the superscript t the point in time; α_0, α_1, β_0, β_1, γ_0 and γ_1 are constant parameters, and ϵ is a random error term. Note that equation (7) is defined for any region i so that, in principle, all six parameters should also carry a superscript i which we leave out to keep the notation as simple as possible.

For estimation purposes, equation (7) can conveniently be rewritten in two alternative versions, namely (i) the generalized error-correction model leading towards a two-step estimation procedure as first proposed by Engle/Granger (EG),[13] and (ii) what may be called the Wickens/Breusch-model (WB) which involves a one-step estimation procedure using instrumental variables.[14]

(i) By some algebraic manipulations of (7), we obtain

$$(8) \quad \Delta \log w_t^i = \beta_0 \Delta \log w_t + \gamma_0 \Delta (z^i-z)_t - \ldots$$
$$\ldots - (1-\alpha_1) [\log w_{t-1}^i - \theta_\alpha - \theta_\beta \log w_{t-1} - \theta_\gamma (z^i-z)_{t-1}] + \epsilon_t$$

with $\theta_\alpha = \alpha_0/(1-\alpha_1)$, $\theta_\beta = (\beta_0+\beta_1)/(1-\alpha_1)$ and $\theta_\gamma = (\gamma_0+\gamma_1)/(1-\alpha_1)$ being the long-run multipliers or co-integration parameters. Comparing the formal structure of (8) and (6), it becomes clear that the term in brackets on the right-hand side of (8) is equivalent to the lagged error term in (6), with the minor difference that (8) has a constant term (θ_α) while (6) has not. Following Engle/Granger, equation (8) can be estimated using a two-step procedure. In the first step, the term in brackets is calculated by running a co-integration regression of the form

$$(9) \quad \log w_t^i = \theta_\alpha + \theta_\beta \log w_t + \theta_\gamma (z^i-z)_t + v_t$$

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13 See Engle, Granger (1987).

14 See Wickens, Breusch (1987).

which yields consistent estimates of the long-run multipliers θ_α, θ_β and θ_γ. Equation (9) can be estimated using ordinary least squares. In a second step, the short-run dynamics is specified in the form of equation (8) with the bracketed term replaced by the lagged error term v_{t-1} from (8), i.e.

$$(10) \quad \Delta \log w_t^i = \beta_0 \Delta \log w_t + \gamma_0 \Delta (z^i-z)_t - (1-\alpha_1) v_{t-1} + \epsilon_t$$

which yields consistent estimates of all relevant short-run coefficients. Again, ordinary least squares can be used. Note that (10) is an equation without a constant term since the long-run constant $\theta_\alpha = \alpha_0/(1-\alpha_1)$ has been estimated in the first step (equation (9)) so that the short-run constant α_0 is identified with the help of the autogressive parameter α_1.

Estimating a co-integration equation like (8) makes only sense if the variable $\log(w_t^i)$ is in fact co-integrated with the linear combination $[\theta_\beta \log w_t + \theta_\gamma (z-z)_t]$. Assuming that both terms are integrated of degree one (I(1)), the estimation of the short-run dynamics will involve only stationary variables. If, in turn, $\log (w_t^i)$ is I(1) and co-integrated with $\log(w_t)$, then - by implication - θ_γ must be zero, i.e. $\gamma_0 = -\gamma_1$. Then the level term (z^i-z) does not appear in the co-integration equation, but the difference term $\Delta(z^i-z)$ is retained in the specification of the short-run dynamics, i.e.

$$(11) \quad \Delta \log w_t^i = \beta_0 \Delta \log w_t + \gamma_0 \Delta (z^i-z)_t - \ldots$$
$$\ldots - (1-\alpha_1) [\log w_{t-1}^i - \theta_\alpha - \theta_\beta \log w_{t-1}] + \epsilon_t.$$

Economically, this means that regional disequilibria may have a short-run impact on the regional wage dynamics which is independent of aggregate disequilibria. However, in the long-run, the regional wage follows the aggregate wage, no matter how severe and sustained the specifically regional disequilibrium has been.[15]

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15 Note that the disequilibirum variable (z^i-z) may be I(0) or I(1), i.e. there may be a stationary difference between the regional and the national labour market disequilibrium or a difference which changes

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(ii) Rewriting (7) in a different form, we obtain

$$(12)\quad \log w_t^i = \theta_\alpha - \phi\, \Delta \log w_t^i + \theta_\beta \log w_t + \theta_\gamma (z^i - z)_t - \ldots$$
$$\ldots - \Pi_\beta\, \Delta \log w_t - \Pi_\gamma\, \Delta (z^i - z)_t + u_t$$

with θ_α, θ_β and θ_γ defined as above, $\phi = \alpha_1/(1-\alpha_1)$, $\Pi_\beta = \beta_1/(1-\alpha_1)$, $\Pi_\gamma = \gamma_1/(1-\alpha_1)$ and $u_t = \epsilon_t/(1-\alpha_1)$. Making the same assumptions on stationarity and/or the degree of integration as above,[16] equation (12) involves the identification of all short- and long-run parameters in a one-step estimation procedure, a result first pointed out by Wickens, Breusch (1987): the long-run multipliers (θ_α, θ_β, θ_γ) are estimated directly, the short-run coefficients (α_0, α_1, β_0, β_1, γ_0, γ_1) are identified by combining the long-run multipliers and the three compound parameters ϕ, Π_β and Π_γ. It is important to note that the error term u_t will be correlated with $\Delta\log(w_t^i)$, which contains a lag-endogenous variable, so that (10) should not be estimated using ordinary least squares (OLS), but rather an instrumental variables technique, with e.g. the lagged right-hand side variables taken as instruments. Hence the benefit of using a one-step procedure as compared to the two-step Engle/Granger approach, which is based on OLS, must be weighed against the cost of losing information in applying instrumental variables instead of ordinary least squares.

As it turned out in most of our estimates, this very information loss is in fact serious enough to make a proper identification of the short-run dynamics virtually impossible - a practical problem which so far seems to have escaped the attention of the recent econometric literature on alternative procedures of estimating short- and long-run parameters. Wickens/Breusch argue that an inadvertent misspecification of the short-run dynamics is most unlikely to harm the estimate of the long-run multipliers and that, for this very reason, there is no point in deleting

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in the long run. In any case, the long-run effect of $[\log w_t^i - \theta_\alpha - \theta_\beta \log w]$ is constrained to zero.

16 I.e. $\log(w_t^i)$ and $\log(w_t)$ be I(1) and $(z_i-z)_t$ be I(0) or I(1).

the short-run dynamics in a first step as done by Engle/Granger.[17] As it stands, this argument is correct since very often, co-integrated variables are highly non-stationary series so that the estimated long-run multipliers will hardly be affected by a lower-order correlation of stationary or close to stationary differences. However, the argument simply bypasses the price to be paid in terms of the information loss which is likely to be important precisely for the identification of the short-run dynamics. After all, the autocorrelation of a variable's intertemporal changes - be they absolute or log-differences (growth rates) - is usually much lower than the autocorrelation of the respective levels so that the standard usage of lagged terms as instruments may involve a quite dramatic loss of efficiency compared to estimates based on ordinary least squares.

This is why we shall mostly focus on the empirical results obtained within the Engle/Granger-framework. As a matter of fact, all estimates of type (7)-equations were made using both procedures alternatively with a fairly clear-cut configuration: whereas the long-run multipliers turned out very similar in the Engle/Granger- and in the Wickens/Breusch-models, the short-run dynamics could only be reasonably identified with the two-step procedure.

As to the definitions of the variables, data limitations commanded some less than perfect choices. The wage is defined as the average wage per hour of a worker in industry as calculated on a representative cross section covering 12 per cent of all industrial establishments with at least ten employees. The wage covers all elements of remuneration which are paid out regularly, i.e. - roughly speaking - the statutory minimum wage plus all fringe benefits which can be assigned to the work in the respective reference period, thus e.g. including overtime or performance premia, but excluding annual gratifications and social security contributions.

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17 See Wickens, Breusch (1987), pp. 35-36

For our purposes, this definition has three pecularities which have to be kept in mind when interpreting the results. First, it is basically a definition of the effective wage, not the contractual minimum wage as fixed in collective agreements. Unfortunately, no comparable data are available on the state level for contractual minimum wages so that no explicit distinction can be made between interregional wage flexibility/rigidity as due to corporatist influences and/or to wage drift.[18] However, as wage drift usually adds an element of flexibility on top of the structure of contractual minimum wages, one may consider any results based on effective wages as indicating an upper limit of flexibility or, conversely, a lower limit of rigidity.

Second, our measure of the wage per hour is not an index calculation, but rather an absolute number (in DM). It is thus not purged of the wage effects of structural change which come about when workers shift from lower to higher paid jobs (or vice versa). While this is a drawback, it should be a tolerable one since our main focus and interest is on interregional wage rigidity, not on intertemporal wage movements. After all, it may not be unrealistic to assume that a measurement error of this kind has little bearing on the interregional structure of industrial wages as the speed and scope of occupational and sectoral labour mobility is likely to depend much more on aggregate cyclical than on specifically regional factors.[19]

Third, as the wage is calculated using a representative cross section of microcensus data, there is likely to be a spurious correlation between the national and any regional wage variable due to an overlapping range of the data base. Given the type of data at our disposal, there is no way

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[18] Regional data on contractual minimum wages are only available in a highly disaggregated form on the level of regional bargaining districts. For technical reasons - in particular the lack of an appropriate weighting scheme to calculate industrial and regional averages -, these data cannot be used for our econometric purposes.

[19] Economy-wide industrial statistics do contain wage <u>indices</u> and we shall discuss estimates with these data further below. By and large, they confirm the conjecture in the text that the measurement error due to structural change is likely to be negligible.

of avoiding this problem by any adjustment procedures simply because no information can be made available on the composition of the representative samples. Naturally, the extent of the problem is likely to depend on the size of the respective region: small states like the cities of Berlin (West), Hamburg, and Bremen as well as Schleswig-Holstein, Rhineland-Palatinate and the Saar may be thought of being small enough to figure as a negligible part of the sample; for large states such as above all Northrhine-Westfalia, Bavaria, Baden-Württemberg, and Lower Saxony, this is not the case. As it turned out, however, our empirical results did not give any clue as to a bias from this source.[20]

The second core variable of our model is (z^i-z), which we have called the specifically regional labour market disequilibrium. The most obvious choice for this variable is the difference (or log-difference) in the unemployment rate between region i and the economy as a whole. Of course, the unemployment rate has its well-known shortcomings as a proxy for a labour market disequilibrium; this is why we also use alternative proxies for (z^i-z) such as the (log-)difference of vacancy rates and the difference of output (employment) ratios between region i and the economy as a whole, with the output (employment) ratio defined as the (log-)deviation of actual (employment) output from a higher-order-deterministic trend in output. Note that the output (employment) ratio is a purely cyclical variable measuring the 'gap' between actual and some sort of natural output (employment) level.[21] As a proxy for a labour market disequilibrium, it catches only cyclical demand variations in the aggregate or in a region (or, for that matter, a sector). Note that the output (employment) ratios are by definition trend stationary; likewise, the difference of any two ratios is stationary so that one can sensibly use it only in the specification (11) which constrains the long-run effect of the disequilibrium variable on the regional wage to zero ($\theta_\gamma = 0$). Also, tests for integration are obiviously redundant for these variables. By and large, the empirical results

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20 Appropriate tests using strictly disjunct sets of data will be presented and discussed below.

21 For the concept of the output ratio, see Gorden (1987), pp. 258-259.

showed our conclusions to be very robust with respect to the proxy used so that we shall mostly confine our subsequent discussion to the model with (z^i-z) being the difference or the log-difference in the respective unemployment rates.

Before actually estimating the EG-model as specified in the equations (8) and (11), we have to scrutinize the univariate time series properties of the main variables concerned. A set of necessary conditions which must be met in order to allow for a meaningful estimation may be summarized as follows:

- for (8): the variables $\log(w^i)$, $\log(w)$ and (z^i-z) are I(1), the first differences $\Delta\log(w^i)$, $\Delta\log(w)$ and $\Delta(z^i-z)$ I(0);

- for (11): the variables $\log(w^i)$ and $\log(w)$ are I(1) and all first differences $\Delta\log(w^i)$, $\Delta\log(w)$ and $\Delta(z^i-z)$ again I(0).

Thus the basic (and only) difference between the two specifications as to the required time series properties is that in (8), (z^i-z) has to be I(1) whereas, in (11), it may be of whatever degree of integration. Hence we first have to make appropriate tests of integration.

In recent years, there has been a dramatic growth of the literature on unit root testing procedures, with the bulk of the publications discussing extensions and modifications of the by now classical Dickey-Fuller test methodology which focuses on the properties of the t-statistics in autoregressions for any variable x_t.[22] Whatever the details of the various tests and their statistical qualities and drawbacks[23], it has

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[22] The most important contributions in this tradition are for the basic model Fuller (1976), and Dickey, Fuller (1979), for appropriate generalizations Said, Dickey (1984), Phillips (1987), and Phillips, Perron (1988). Most recently, Schmidt, Phillips (1989) have established a new strand of unit root tests which look 'directly' at the asymptotic behavior of a series x_t (see also Campbell, Perron (1991), p. 17). In the following analysis, we shall remain in the realm of the 'traditional' DF-type models.

[23] For a survey evaluation, see Campbell, Perron (1991), pp. 8-27.

become more and more evident that, for the applied econometrician, the outstanding difficulty of all tests is to discriminate between a stochastic and a deterministic trend component in the data generation process. Following the seminal work of Nelson, Plosser (1982), who used simple DF-tests for a wide range of macroeconomic variables, the maintained hypothesis was that most time series follow a unit root process. However, more recent theoretical contributions to the debate - notably Blough (1988), Cochrane (1991) and Campbell, Perron (1991) - have forcefully argued that there is a near observational equivalence of trend and difference stationary processes which may seriously impair the power of the tests if only the deterministic trend is adequately specified, i.e. possibly including non-linear components. In practice, the simple experimentation with higher order deterministic trend terms regularly shows that many time series can be consistently interpreted in two alternative ways, either as a difference stationary process or as a trend stationary process with the trend usually being of a higher than first order. Hence, very much care must be taken in specifying the deterministic trend to be imposed on the data, in any case much more care than in the early unit root testing literature following and including Nelson, Plosser (1982).

Given the paramount importance of deterministic trend specification, we shall adopt a very broad testing procedure involving five augmented Dickey-Fuller (ADF)-tests which differ in the choice of trend components. The basic ADF-equation for any variable x_t is given by

$$(13) \quad \Delta x_t = -(1-\beta)\, x_{t-1} + \gamma\, \Delta x_{t-1} + \epsilon_t,$$

with $-(1-\beta)$ and γ defined as the coefficients to be estimated and ϵ_t as a random error term. If $-(1-\beta)$ turns out to be negative at a standard significance level as given by the appropriate DF-test-tables, then I(1) can be rejected; we shall denote the t-ratio of the parameter $-(1-\beta)$ as DF_0. Note that - following standard practice - the 'lagged endogenous variable' Δx_{t-1} will be included on purely pragmatic grounds if a first regression of Δx_t on x_{t-1} gives a residual with a substantial autocorrelation; if not, a simple 'non-augmented' DF-test (with $\gamma=0$) is used.[24]

Equation (13) sets the null hypothesis of I(1) against stationarity without any allowance for a deterministic trend component, not even a drift term or a linear trend. To offer a broad spectrum of alternative test specifications with other, more general assumptions on the deterministic trend, we then apply four different variants of a two-step DF-test, with the first step designed so as to purge the series x_t of the relevant deterministic components. Hence, a regression of the form

$$(14)\quad x_t = \sum_{i=0}^{n} \alpha_i t^i + y_t$$

is run for i = 0, 1, 2, 3, with t defined as a trend variable and y_t as the respective 'detrended' series of x_t, i.e. that part of the variation of x_t which cannot be accounted for by deterministic elements. Thus, step by step, higher-order deterministic trends are introduced, namely a constant (α_0), a linear ($\alpha_1 t$), a quadratic ($\alpha_2 t^2$) and a cubic trend term ($\alpha_3 t^3$). With the detrended variable y_t, the ADF-test regression

$$(15)\quad \Delta y_t = -(1-\beta')\, y_{t-1} + \gamma' \Delta y_{t-1} + \epsilon_t'$$

is run, with $-(1-\beta')$ and γ' as the coefficients to be estimated and ϵ_t' the respective random error term. The t-ratio of $-(1-\beta')$ gives the appropriate DF-statistics, which we call DF_1, DF_2, DF_3 and DF_4 respectively. In tests for integration of degree two, exactly the same procedures are used to calculate DF_0 down to DF_4, with the only difference being that x_t is replaced by Δx_t in equations (13) and (14).[25]

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[24] For a survey on the question of augmentation, which is more important for data with seasonal components, see Campbell, Perron (1991), pp. 14-16.

[25] Note that, only for DF_1 and DF_2, the two-step procedure as described in the text is asymptotically equivalent to the one-step procedure with the trend variable simply included in the regression of ...

Our testing design gives a broad menue of analogously constructed test statistics for different deterministic trend constellations. It thus allows to detect the breaking point of deterministic complexity - if there is one - at which a non-stationary series can be described as being stationary around some possibly rather complex long-term trend. This is very helpful for two rather pragmatic reasons. First, a casual glance over the relevant time series - for that matter, over virtually all macroeconomic time series for the post-World War II period - indicates that, if at all, a second- or third-order deterministic trend - and not a linear one - could be made responsible for the observed non-stationarity. To stop the testing at a linear trend may thus simply bypass an important prima facie characteristic of the series. Second, as the integration test is no more than the preliminary stage for subsequently estimating co-integration equations, it is important to know whether the relevant two series can be described not only as difference stationary, but also as trend-stationary with the same order of the trend terms. If this is the case, then the co-integration equation can be interpreted as a test for deterministic co-integration - meaning that the same vector of parameters, which describes the unit root, also eliminates the deterministic trend from the data. Note that deterministic integration is an unambiguously stronger demand than mere 'stochastic' integration, the latter meaning that the relevant parameter vector removes the unit root, but not the deterministic trend, which has to be eliminated before running the co-integration regression.[26] Statistically, deterministic co-integration is, of course, a very convenient property since it spares the effort of filtering out deterministic components, which, in all likelihood, are close to observationally equivalent to a unit root process. Economically, deterministic co-integration has the advantage of greatly facilitating the interpretation of co-integration as a long-run equilibrium: whatever the data generation process of the relevant time series may be - a unit root

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type (13); for DF_3 and DF_4, which concern higher order than linear trends, the two-step and the one-step procedure may not be equivalent, even not asymptotically, with the standard properties of the tests applying only in the two-step version. See Campbell, Perron (1991), pp. 8-12.

26 For the details of the two concepts, see Campbell, Perron (1991), p. 25.

or a complex deterministic trend -, the series can be taken to have an unambiguous long-run equilibrium relationship.

It is important to realize that our testing methodology as described above reflects a rather pragmatic philosophy which considers the DF-test battery more as a set of tools to skan the data for the characteristics of their generation process, not to obtain a clear-cut answer from some nicely tabulated test statistics on how the data generation process is likely to be. As experience shows, all standard tests loose much of their power if higher-order deterministic trends are allowed for. Hence, one should simply not expect anything like a straight and unmistakable answer from them, at least not in the world of the sample sizes we have at hand in economic research. Nevertheless, the skanning of the data for trend components and unit root processes is extremely important so as to lay the ground for the subsequent procedures to be applied, notably in multivariate co-integration analysis.

In our view, the data skanning can reasonably be made with many different test types, be they of the Dickey-Fuller or, e.g. of the Schmidt-Phillips variety.[27] Given the uniformly low power of these tests when putting difference against higher-order trend stationarity, the choice of the test family becomes a marginal issue compared to the appropriate specification of the deterministic trend. After all, the low test power is the consequence of the near observational equivalence of the alternative hypotheses and the low informational content of the data; the first is a theoretical problem beyond the reach of statistical medicine, the second could be cured if richer data were available. Note that 'richer data' does of course not simply mean more data: e.g., if the relevant long-term for integration and co-integration tests is to be defined in decades, then a mere increase of the data frequency by switching from annual to quarterly or monthly data, if available, cannot really help, because the high frequency range is likely to have very little marginal information on long-run univariate movements or multivariate relationships. To the contrary, the test power is usually lower for data

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27 See Schmidt, Phillips (1989).

of frequencies shorter than a year as seasonal adjustment procedures must be applied and often lead to less precise and biased estimates.[28]

4.2. Estimation Results

Let us first discuss the main results of our integration tests. Table 1 presents the five test statistics DF_0, DF_1, DF_2, DF_3, and DF_4 for the variables $\log(w^i)$, $\log(u^i/u)$, and (u^i-u), both for first- and for second-order integration. As the period covered is 1950-89, the Saar and Berlin remain exluded since not all of the relevant data are available for these two states in the first decade of the sample period. The DF-tests were made in the augmented form involving one lag-endogenous term for first-order tests and in the non-augmented form ($\gamma=0$) for the second-order tests.[29] Note that critical values at the standard significance levels are only available for the statistics DF_0, DF_1, and DF_2,[30] but not for DF_3 and DF_4. Nevertheless, the magnitude of the statistics DF_3 and DF_4 relative to DF_0, DF_1, and DF_2 gives us a clear indication of whether the additional trend terms are of much help to track the long-run movement of the time series. This is in line with our pragmatic philosophy on testing sketched above.

The picture conveyed by the numbers in Table 1 is informative though as usual not conclusive. As to the wage variable ($\log(w^i)$ and $\log(w)$), the 'low-order' DF-statistics (DF_0, DF_1, DF_2) uniformly indicate that integration of degree one (I(1)) cannot be rejected at any of the standard significance levels (1 to 10 per cent). However, the picture changes when higher-order trend terms are added. In particular, the inclusion of a cubic trend term leads to a dramatic increase of the

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28 On this issue, see Shiller, Perron (1985), Ghysels, Perron (1990), Jaeger, Kunst (1990), and the survey in Campbell, Perron (1991), pp. 13-14.

29 The basic pattern of the results is hardly affected by the pragmatic decisions on restricting or not the parameter γ in the equations (13) and (15). The choice of augmentation or non-augmentation is made on the basis of the autocorrelation in the error term of the non-augmented test version.

30 See Fuller (1976), p. 373.

Table 1 - Tests for Integration, West German States 1950-1989

Variable:		$\log w^i$		$\log(u^i/u)$		u^i-u	
	Order:	1st	2nd	1st	2nd	1st	2nd
Schleswig-Holstein	DF_0	1.94	-2.31	-2.01	-5.44	-8.23	-2.71
	DF_1	-0.59	-5.23	-1.73	-5.64	-5.46	-2.83
	DF_2	-0.39	-5.85	-2.39	-5.77	-3.50	-3.74
	DF_3	-1.67	-6.63	-3.94	-5.82	-3.62	-5.80
	DF_4	-3.66	-6.63	-5.21	-5.83	-4.05	-7.23
Hamburg	DF_0	1.34	-1.64	-2.35	-3.70	-3.25	-3.49
	DF_1	-0.69	-3.83	-2.35	-3.69	-3.35	-3.48
	DF_2	-0.53	-4.31	-2.18	-3.93	-3.22	-4.09
	DF_3	-1.66	-5.41	-5.40	-3.86	-4.61	-3.93
	DF_4	-4.31	-5.41	-5.30	-4.06	-4.62	-3.68
Lower Saxony	DF_0	1.20	-1.62	-1.59	-6.19	-3.54	-5.05
	DF_1	-0.33	-3.46	-3.38	-6.24	-3.90	-5.17
	DF_2	-0.69	-3.89	-3.22	-6.40	-3.81	-6.79
	DF_3	-2.13	-4.72	-3.38	-6.44	-3.66	-9.02
	DF_4	-4.74	-4.64	-3.33	-6.44	-3.56	-8.87
Bremen	DF_0	1.63	-1.91	-0.70	-6.68	0.32	-6.43
	DF_1	-0.27	-4.25	-2.94	-6.68	-0.17	-6.42
	DF_2	-0.79	-4.48	-3.00	-6.67	-1.71	-7.57
	DF_3	-2.07	-5.44	-2.93	-6.77	-3.18	-7.21
	DF_4	-4.97	-5.32	-4.05	-6.75	-3.36	-6.99
Northrhine-Westfalia	DF_0	1.74	-2.12	-2.22	-4.82	-2.39	-2.70
	DF_1	-0.10	-4.19	-2.32	-5.21	-2.67	-3.20
	DF_2	-0.69	-4.48	-1.83	-5.66	-3.38	-3.79
	DF_3	-2.78	-5.10	-2.48	-5.74	-2.95	-4.64
	DF_4	-4.64	-4.91	-3.68	-5.86	-2.24	-6.50
Hesse	DF_0	1.57	-1.95	-1.65	-4.86	-1.87	-3.34
	DF_1	-0.05	-3.69	-3.48	-4.86	-2.34	-3.36
	DF_2	-1.19	-3.74	-3.58	-4.88	-2.98	-3.94
	DF_3	-2.90	-4.94	-3.73	-4.88	-3.95	-4.98
	DF_4	-5.00	-4.77	-4.44	-4.88	-3.83	-4.94
Rhineland-Palatinate	DF_0	1.33	-1.79	-2.18	-7.06	-6.53	-5.63
	DF_1	-0.27	-3.51	-2.46	-7.14	-6.10	-5.92
	DF_2	-1.21	-3.70	-3.66	-7.22	-3.96	-8.34
	DF_3	-2.58	-4.39	-5.81	-7.15	-4.16	-9.38
	DF_4	-4.77	-4.28	-7.86	-7.05	-4.42	-8.98
Baden-Württemberg	DF_0	1.64	-1.86	-1.01	-5.74	-2.19	-2.77
	DF_1	-0.26	-4.17	-1.20	-5.80	-2.80	-2.82
	DF_2	-0.86	-4.33	-2.13	-5.81	-2.81	-3.64
	DF_3	-2.14	-5.70	-2.01	-5.99	-2.37	-5.44
	DF_4	-3.84	-5.62	-3.48	-6.02	-3.29	-5.32
Bavaria	DF_0	1.12	-1.41	-0.43	-5.48	0.15	-3.32
	DF_1	-0.44	-3.34	-1.14	-5.59	-0.52	-3.55
	DF_2	-1.22	-3.60	-4.14	-5.70	-2.10	-3.86
	DF_3	-2.16	-4.28	-4.93	-5.70	-3.09	-3.92
	DF_4	-4.46	-4.26	-4.90	-5.73	-4.22	-4.02
West Germany	DF_0	1.34	-1.70	-	-	-	-
	DF_1	-0.27	-3.55	-	-	-	-
	DF_2	-0.96	-3.85	-	-	-	-
	DF_3	-2.50	-4.56	-	-	-	-
	DF_4	-4.72	-4.47	-	-	-	-

Notes: Test statistics DF_0, DF_1, DF_2, DF_3 and DF_4 as defined in the text, with (non-augmented) Dickey-Fuller test ($\gamma, \gamma' = 0$ in equations (13) and (15)) for first-order tests and augmented Dickey-Fuller test (unrestricted γ, γ' in equations (13) and (15)) for second-order tests. Critical values for rejecting the null hypothesis of a unit root at the 5 %- (10 %-) level are about -1.95 (-1.61) for DF_0, -2.93 (-2.60) for DF_1, and -3.50 (-3.18) for DF_2 (see Fuller 1976, p. 373, Table 8.5.2, for n=50). For DF_3 and DF_4 no such critical values are available, but it is certain that they will be higher in absolute terms than the ones for DF_2.

DF-statistic, with DF_4 being well above four in absolute terms in eight out of nine cases. The corresponding autoregressive parameter β (not printed in the table!) turned out to be lower than 0.5 in most cases. For our prospective co-integration analysis, it is further important to realize that the structure of the test statistics is very similar for all $\log(w^i)$ and for $\log(w)$ which can all be taken either to follow a unit root-process or to be stationary around a third-order deterministic trend. Hence, a subsequent co-integration estimate can be quite safely interpreted as a test for the stronger form of deterministic, not just the weaker form of stochastic integration.

As to second-order integration of the wage variables (in logs), the DF-statistics speak a fairly unambiguous language: all of them point against I(2), as far as they go at extremely high significance levels, and higher-order deterministic trends beyond a constant and a linear trend do hardly matter. Again, the pattern of DF-statistics between states and for West Germany as a whole is fairly uniform so that, again, the same data generation process can be assumed to drive the wage in all regional parts of the country.

As to the two variables for the regional labour market disequlibrium ($\log(u^i/u)$, u^i-u), there is no doubt that I(2) cannot be accepted since, in the vast majority of cases, the DF-statistics turn out to be very high in absolute terms, and if measurable, highly significant. Also, the detrending does not affect the results to any substantial degree which points to the absence of strong linear or non-linear trend elements. On the other hand, the results for the I(1)-tests differ markedly between states, between the disequilibrium variables used and between the type of detrending performed, with no easy generalizations possible. In some states like Schleswig-Holstein, Hamburg and Rhineland-Palatinate, the test statistics point more towards stationarity although - even for them - the results are not altogether clear-cut. For other states like Bremen, Baden-Württemberg and Bavaria, the weight of the testing evidence pulls towards a unit root. For some of the states like Schleswig-Holstein and Rhineland-Palatinate, it also strongly depends on the variable used. All in all, it seems that both assumptions (I(0) and I(1)) can be defended on statistical grounds at least for a subset of states. Hence, in view of $\log(w^i)$ and $\log(w)$ being I(1), none of the specifications (8) and (11) can be excluded a priori on basis of univariate time series characteristics of the data so that we may start our empirical exercise with either of them. For reasons of convenience, we shall first pick the more restrictive version (11) so as to find out whether the aggregate wage level 'alone' delivers a satisfactory account of the regional wage levels in the long run, i.e. whether log (w^i) and log (w) are co-integrated.

Let us now turn to the results of our estimates of the co-integration equation. To begin with, Table 2 presents the coefficients of a two-step Engle/Granger-model of the type (11), i.e. co-integration is tested just between the regional and the aggregate wage without consideration for specifically regional labour market disequilibria in the long run. All across the nine states of the sample, the estimates of Table 2 have some characteristic features which we shall briefly summarize and discuss in the following paragraphs.

(i) As to the statistical efficiency of the long-run estimates, the two coefficients θ_α and θ_β appear to be very well identified in all equations of the first step of the EG-procedure: the lowest $\bar{R}^2$ is still as high as 0.9992, the standard estimation errors are extremely low and the t-ratios of θ_β fall in the range 200-500 which is large by any reasonable standard.

(ii) As to the magnitude of the long-run coefficients, a clear-cut pattern can be recognized. For all states, θ_β is rather close to unity; in most cases, however, it remains significantly different from one since the very low standard estimation errors make the confidence intervals extremely small and thus allow to reject the null hypothesis that θ_β exactly equals unity. In all equations, θ_β falls within the interval [0.98; 1.03], which is narrow indeed; it means, e.g., that a 10 %-wage increase in West Germany as a whole would be accompanied by a 9.8 %-increase in the 'lower limit'-state Northrhine-Westfalia and a 10.3 %-increase in the 'upper limit'-state Rhineland-Palatinate, with all other states falling in between. To put it in historical terms: if the two 'limit'-states had started in 1950 at the economy-wide average wage of 1.22 DM per hour (which they have not!) and if the West German economy had experienced its actual average wage growth of 7.24 % p.a. - leading to an average wage of 18.63 DM per hour by 1989 -, the 'lower-limit'-state would have

Table 2 - Estimated Coefficients of the Engle/Granger-Model for West German States 1950-1989

	long-run (1st step)							short-run (2nd step)					short-run alternatively		
	const	log w	$\bar{R}^2$	SE(%)	DW	DF	ρ	Δ log w	$\Delta \log(u^i/u)$	v_{-1}	$\bar{R}^2$	DW	$\Delta(u^i-u)$	$\Delta(OR^i-OR)$	$\Delta(ER^i-ER)$
Schleswig-Holstein	-0.0584	1.0199	0.9992	2.44	1.2585	-4.0813	0.3628	1.0237	-0.0031	-0.7091	0.6684	1.9715	0.0040	0.0328	0.3002
	(0.0086)	(0.0045)						(0.0557)	(0.0048)	(0.1776)			(0.0077)	(0.2410)	(0.2976)
Hamburg	0.0851	1.0125	0.9994	2.10	0.5970	-3.4640	0.6102	0.9847	0.0288	-0.5008	0.8256	1.5982	0.0030	-0.0117	-0.0714
	(0.0073)	(0.0039)						(0.0303)	(0.0196)	(0.1318)			(0.0040)	(0.1656)	(0.1479)
Lower Saxony	-0.0317	1.0190	0.9998	1.20	0.9312	-3.3520	0.5326	1.0054	-0.0008	-0.4920	0.8860	2.0465	-0.0010	-0.1894	0.0276
	(0.0042)	(0.0022)						(0.0227)	(0.0019)	(0.1564)			(0.0051)	(0.1685)	(0.1970)
Bremen	0.0398	0.9998	0.9993	2.36	0.3755	-1.9813	0.8093	0.9758	-0.0046	-0.2105	0.8076	1.6849	0.0026	-0.1819	0.1412
	(0.0082)	(0.0043)						(0.0306)	(0.0128)	(0.1108)			(0.0036)	(0.1320)	(0.0929)
Northrhine-Westfalia	0.0717	0.9828	0.9998	1.27	0.5600	-4.3774	0.5635	1.0076	0.0021	-0.4953	0.9470	1.9241	-0.0086	-0.1825	0.5271
	(0.0045)	(0.0023)						(0.0174)	(0.0015)	(0.1150)			(0.0055)	(0.1671)	(0.2871)
Hesse	-0.0038	1.0045	0.9995	2.04	0.3524	-2.3942	0.7850	0.9953	-0.0343	-0.1775	0.8743	1.5701	-0.0239	0.4229	0.4077
	(0.0071)	(0.0038)						(0.0251)	(0.0217)	(0.1028)			(0.0109)	(0.2063)	(0.2802)
Rhineland-Palatinate	-0.0964	1.0283	0.9997	1.54	0.6034	-2.7104	0.6873	1.0305	-0.0145	-0.3077	0.8928	1.8280	-0.0025	0.3564	-0.1180
	(0.0054)	(0.0028)						(0.0248)	(0.0135)	(0.1298)			(0.0066)	(0.1301)	(0.1695)
Baden-Württemberg	-0.0600	1.0225	0.9994	2.21	0.4329	-3.4862	0.6752	0.9944	-0.0055	-0.3547	0.8462	2.1856	-0.0037	0.2507	0.1225
	(0.0077)	(0.0041)						(0.0285)	(0.0134)	(0.1025)			(0.0069)	(0.2619)	(0.2588)
Bavaria	-0.1307	1.0194	0.9995	1.96	0.5614	-3.4140	0.6344	0.9841	-0.0236	-0.3756	0.8311	1.8288	-0.0085	-0.3065	0.1802
	(0.0069)	(0.0036)						(0.0280)	(0.0295)	(0.1202)			(0.0127)	(0.3270)	(0.3011)

Notes: $\bar{R}^2$ = adjusted R^2; SE = standard estimation error; DW = Durbin-Watson-statistic; DF = Dickey-Fuller-statistic for first-order autoregression of residuals; ρ = estimated coefficient of first-order autoregression of residuals; v_{-1} = lagged error term from first step of EG-model. Table based on estimates of equation (11); first step specified by equation (9) with θ_r = 0, second step by equation (10). Standard estimation errors of coefficients in parenthesis. For DW, the critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about 0.755 (range of indifference from 0.484 to 0.755); see Sargan, Bhargava (1983), p. 157, Table I for T=51 and n=1. For DF, critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about -3.365, at the 10 %-level about -3.066 and at the 15 %-level about -2.864; see Phillips, Ouliaris (1990), p. 190, Table II for n=1.

ended up at a wage of 18.33 DM per hour, the 'upper-limit'-state at 19.12 DM per hour, a less than 1 DM-difference after 39 years!

The long-run constant θ_a varies depending on whether, roughly speaking, the respective state started as a high-wage or a low-wage region. E.g., the northern states Northrhine-Westfalia, Hamburg, and to a lesser extent Bremen had traditionally high industrial wages while the southern states Rhineland-Palatinate, Baden-Württemberg and especially Bavaria lagged behind. Remarkably enough, it is the originally low-wage states of Southern Germany which experienced the most rapid trend growth of wages and thus somewhat caught up to the northern industrial centers. Up to the present, this catching-up process has transformed Baden-Württemberg, but not yet Bavaria and Rhineland- Palatinate into high-wage regions: e.g., by 1989, the original 15 %-wage lead of Northrhine-Westfalia over Rhineland-Palatinate has been narrowed (but not removed!) to 3.5 %; in turn, the lag of Bavaria behind Northrhine-Westfalia has shrunk from about 20 % in 1950 to less than 7 % by 1989, but still it is a lag.[31] At this point, it is important to remember that, within the confines of our model, these catching-up processes have nothing to do with labour market disequilibria in the respective regions; rather they reflect changing industrial structures between regions and/or changes in the long-term wage positions of those industries which are disproportionally represented in the particular region in question. If, e.g., the wage lead of Northrhine-Westfalia has been gradually eroded, this means that the workers in typically capital-intensive industries which are concentrated in the Rhine/Ruhr-valley, have experienced a long-run terms-of-trade loss vis-à-vis workers in other industries which dominate in other regions. There are many plausible reasons for such long-term shifts as, e.g., a changing interindustrial pattern of qualification, of age composition, of the share of male and female workers etc.; per se, however, such shifts are independent of the extent of regional labour market disequilibria, be

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31 Note that these numbers have been calculated on basis of ex-post predictions of regional wages given the long-run coefficients presented in Table 2. However, if calculated on basis of actual wages, the numbers are not much different.

they temporary or permanent. In any case, they are not the focal point of our empirical analysis.[32]

(iii) In testing for co-integration, we use two test-statistics, namely DW - the Durbin-Watson statistic following the rationale of the so-called Sargan-Bhargava-test[33] - and DF, the conventional Dickey-Fuller statistic calculated by imposing a pure first-order autocorrelation process (without drift or trend terms) on the residuals of the co-integration equation.[34] Note that the critical values for both statistics if used in tests for co-integration must be based on the case of a random walk with drift, i.e. an autoregressive process involving a constant, because the co-integration equation itself does have a constant term (θ_α).[35] This, in turn, is the reason why the power of the Sargan-Bhargava-test suffers from a wide margin of inconclusiveness where the decision on accepting or rejecting the unit root process cannot be made at any of the standard confidence levels; e.g., in the case of 51 observations and a 5 %-confidence level, the 'interval of indifference' for the DW-statistic is [0.484; 0.755], in practice quite a large range as will become clear below. Therefore we shall mostly focus on the message of the DF-statistic.

In glancing over the actual co-integration test results of Table 2, the picture is ambiguous: the DW- and DF-statistics point to a considerable degree of autocorrelation of the residuals in the first step of the EG-procedure, but whether the null hypothesis of a unit root should be rejected at standard significance levels is unclear. The DW-statistic

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32 As a matter of fact, they would have to be explained in a completely different conceptual framework aimed at exploring the determinants of the long-run wage structure. See, e.g., Dickens, Katz (1987a, b), Gibbons, Katz (1989), Katz, Summers (1989), Krueger, Summers (1988), Thaler (1989), and Gundlach, Fels (1990).

33 See Sargan, Bhargava (1983).

34 Unlike in tests for integration, it makes no sense to include deterministic elements in the DF-test simply because, by construction, an error term of an ordinary least squares estimate has a zero mean.

35 The relevant critical values for an infinite sample size are tabulated in Phillips, Ouliaris (1990), pp. 189-190.

indicates as much as four cases of indifference, and the DF-statistic rejects the null hypothesis in five out of nine equations at the 5 %-level and is very close to rejection in a sixth case. Note also that the autoregressive parameter ρ is far below one in most of the estimates underlying the DF-statistics - in seven out of nine cases below 0.7; nevertheless, standard errors in the range of 0.10-0.15 do often not allow to reject the unit root-hypothesis at the conventional levels of significance. As usual, it seems to be virtually impossible to discriminate between a stationary process with autoregressive parameter close to but below one, and a random walk.[36]

In view of the extremely tight statistical relationship between $\log(w^i)$ and $\log(w)$ as measured by $\bar{R}^2$, however, it would look somewhat farfetched to decide against co-integration only because the null hypothesis of a unit root cannot be rejected at the usual significance levels; after all, these very significance levels are quite demanding for the particular test purpose at hand. As Hendry (1986)[37] has pointed out, the likelihood of a type II-error - falsely accepting a wrong null hypothesis - may become very high if the alternative hypothesis of stationarity includes cases with autoregressive parameters of close to one. Given the rather strong a priori belief in the existence of a long-run equilibrium relationship between the wage in region i ($i = 1, 2, \ldots, n$) and the country as a whole, the price to be paid for any such type II-error would also be very high; it would mean rejecting co-integration despite strong priors in its favour and a very high $\bar{R}^2$. A careful specification of an appropriate loss function (of course, going way beyond the scope of our analysis) would certainly give due weight to the priors and thus call for less demanding significance levels.

Looking over the test statistics DW and DF with these second thoughts in mind, the general conclusion should be that a substantial autocorrelation

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36 This, again, is a conseqence of the near observational equivalence of a unit root process and a stationary or trend-stationary process. See Blough (1988), Cochrane (1991), and Campbell, Perron (1991), pp. 18-23.

37 See Hendry (1986), p. 206.

in residuals is discernible, but that it does not quite point towards a unit root. This conjecture derives some support from the results of the second step of the EG-procedure which yields an autoregressive parameter α_1 below 0.7 in seven out of nine cases - thus indicating that the short-run dynamics does not involve anything close to a unit root process. Again, however, the respective standard errors - they are in the range of 0.10-0.15 - do not quite allow to reject the unit root-hypothesis at the all too demanding conventional levels of significance.

(iv) The very high $\bar{R}^2$ all across the co-integration equations of Table 2 raises the question of how much 'autonomous' variation of the regional wage is really left for the short-run dynamics to be explained in the second EG-step. In a log-linear specification, the standard error is also a good measure of the 'standard percentage gap' between the actual and the fitted values so that it can serve as a proxy for the 'leeway' which is left for the second step. As the standard errors given in Table 2 show, this leeway is quite small all throughout, generally in the range of 1-2.5 %. Thus, before proceeding any further, it is important to keep in mind that the 'explanatory scope' left for any specifically regional factors is quite small since more than 97 % of all variation of regional wages can be accounted for by the movement of the aggregate wage. Hence, whatever measure of regional wage flexibility we shall obtain, it will describe no more than a marginal phenomena.

(v) The short-run dynamics in Table 2 is a bit more differentiated across states than the long-run relationship although some important common features can again be recognized. First, all coefficients β_0 come close to one at very low standard estimation errors; they all fall in the narrow interval [0.97; 1.03] with none of them being different from unity at standard significance levels. Hence it is quite safe to conclude that any regional wage and the aggregate wage share not only a common stochastic or deterministic trend, but also a common short-run dynamics. Second, all coefficients of the variable $\Delta\log(u^i-u)$ which serves as a proxy for the rate of change in the specifically regional labour market disequilibrium, turned out to be not different from zero at the conventional significance levels, although - in seven out of nine cases - they have the expected negative sign. Third, the autoregressive parameter α_1 of the lagged error term from the co-integration equation varies considerably across states, just as the co-integration test statistics DW and DF vary in the first step of the EG-procedure. However, α_1 falls in the central range 0.3-0.7 in seven out of nine cases at standard errors around 0.10-0.15 - thus in general pointing towards a strongly lagged adjustment, but not a unit root process.[38] Finally, the statistical quality of the short-run estimates is quite satisfactory, except perhaps for the equation of Schleswig-Holstein, which has a much lower explanatory power than all other estimates; in general, the $\bar{R}^2$ is well above 0.8 and the DW-statistic does not indicate any substantial autocorrelation in the residuals.[39]

(vi) As the parametric shape of the short-run dynamics might be very sensitive with respect to the disequilibrium variable used, we also carried out the second step of the EG-procedure using three alternative proxies for $\Delta(z^i-z)$ in equation (11), namely the change of the absolute difference of the regional and the aggregate unemployment rate $(\Delta(u^i-u))$,[40] the change of the difference of the output (employment) ratio between region i and the economy as a whole $\Delta(OR^i-OR)$, $\Delta(ER^i-ER)$ respectively), with the ouput (employment) ratio again defined as the log-deviation of the actual output (employment) from a higher-order deterministic trend. Remember that the output and the employment ratios need not be tested for their degree of integration since, by definition, they are constrained to be trend stationary. Table 2 presents just the coefficient γ_0 of these variables in the relevant regressions; the concommittant parameters β_0 and $(1-\alpha_1)$ did hardly differ from previous

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38 In general, α_1 comes very close in magnitude to the autoregressive parameter ρ underlying the DF-statistics in the first step of the EG-estimate.

39 As the DW-statistic is biased in the presence of lag-endogenous variables, alternative autocorrelation statistics like the Durbin-t-statistic were also applied. They did not indicate any residual autocorrelation either.

40 Note that, in an alternative set of specifications, we also replaced the unemployment rate by the vacancy rate as a proxy for the labour market disequilibrium, both in the specifications with $\Delta\log(u^i-u)$ and $\Delta(u^i-u)$. As a matter of fact, none of our conclusions in the text was substantially affected by this change.

estimates so we simply left them out of the table. Qualitatively, the results confirm our prior conjectures: for $\Delta(u^i-u)$, β_0 has the expected negative sign in six out of nine equations, but only in one of them, the one for Hesse, it is also significantly different from zero at the 5 %-level. The record of $\Delta(OR^i-OR)$ and $\Delta(ER^i-ER)$ as proxies is equally poor: in the case of $\Delta(OR^i-OR)$, β_0 has the expected positive sign only in four out of nine equations, with significance at the 5 %-level being achieved only in two (Rhineland-Palatinate and, again, Hesse); in the case of $\Delta(ER^i-ER)$, there is seven times the expected positive sign, but in just one case (Northrhine-Westfalia) at the 5 % significance level.

To gain a rough quantitative impression of the magnitude of the short-run wage effects initiated by specifically regional labour market disequilibria, let us assume that $u = u_{-1} = u_{-1}^i$ and $u^i = 2u_{-1}^i$, i.e., from an initial state of equal regional and aggregate unemployment rates - say 1 or 4 % -, the regional rate doubles. Table 3 summarizes the one-period impulse effect of this exogenous shock on the regional wage for all nine states, given the parametric shape of the EG-estimates in Table 2. Note that, due to the restriction $\gamma_0 = -\gamma_1$ which is imposed in equation (11), the impulse effect in the first period is also the maximum effect to be reached at all since, from the second period on, it gradually dissipates with the speed of the decay depending on the magnitude of the autoregressive parameter α_1.

As can be seen from the numbers in the table, the impulse effects are small by any standard, if at all 'correctly' signed: in the log-linear framework, a doubling of the unemployment rate differential leads to a temporary dampening of regional wage growth by more than 1 % in just three states, by more than 2 % in just one state (Hesse), on average by just 0.43 %. The semi-log model implies a somewhat greater wage effect in higher unemployment ranges - on average a dampening by 2.14 % at a 5 % aggregate unemployment rate -, but that should not affect the main thrust of our conclusions; after all, a doubling of the unemployment rate differential under these circumstances would be an altogether spectacular exogenous shock, which is way off the magnitude of any actual event for the relevant states in the sample period. In this sense, even a 12 %-regional wage effect as the one indicated by the estimates for Hesse may still look quite modest, not to mention the 4-4.5 %-effect for

Table 3 - Impulse Effect on the Regional Wage (in %) of a Doubling in the Regional Labour Market Disequilibrium

$u_{-1} = u^i_{-1} = \ldots$	Log-linear 0-100 %	Semilog-linear Model 1 %	5 %
Schleswig-Holstein	-0.21	+0.40	+2.00
Hamburg	+2.00	+0.30	+1.50
Lower Saxony	-0.06	-0.10	-0.50
Bremen	-0.32	+0.26	+1.30
Northrhine-Westfalia	+0.15	-0.86	-4.30
Hesse	-2.38	-2.39	-11.95
Rhineland-Palatinate	-1.01	-0.25	-1.25
Baden-Württemberg	-0.38	-0.37	-1.85
Bavaria	-1.64	-0.85	-4.25
Arithmetic Average	-0.43	-0.43	-2.14

Notes: Impulse effects calculated on basis of the respective estimates in Table 2.

Northrhine-Westfalia and Bavaria and the less than 2 %-effect for Lower Saxony, Rhineland-Palatinate and Baden-Württemberg. On the whole, the picture of the short-run wage dynamics is clearly one of a high degree of rigidity with respect to regional labour market disequilibria.

To test for intertemporal stability of the parameters in Table 2 and to include the two additional states Berlin and the Saar, we reestimated equation (11) and the respective test statistics for the shorter period 1960-89. The results of the integration tests are summarized in Table 4. By and large, they support our prior conjectures as to the degree of integration of the three variables concerned: $\log(w^i)$ and $\log(w)$ are either difference stationary (I(1)) or trend stationary around a higher order (notably a cubic) deterministic trend; $\log(u^i/u)$ is either I(0) or I(1), again depending on the complexity of the deterministic trend, in any case not I(2). Only the time path of (u^i-u) looks slightly different from Table 2 where it appeared to be I(0) or I(1); in Table 3, the test statistics point much more towards I(1), even if higher order deterministic trend terms are allowed for.

Table 4 - Tests for Integration, West German States 1960-1989

	Variable:	$\log w^i$		$\log(u^i/u)$		u^i-u	
	Order:	1st	2nd	1st	2nd	1st	2nd
Schleswig-Holstein	DF_0	1.89	-2.15	-2.10	-5.28	-0.22	-6.16
	DF_1	-0.34	-3.90	-2.09	-5.33	-1.05	-6.22
	DF_2	-0.51	-4.70	-1.97	-5.45	-2.21	-6.38
	DF_3	-2.48	-5.03	-4.11	-5.30	-3.90	-6.41
	DF_4	-2.94	-5.07	-4.21	-5.15	-3.86	-6.47
Hamburg	DF_0	1.44	-1.72	-1.34	-3.46	-1.71	-2.08
	DF_1	-0.54	-3.18	-1.33	-3.45	-1.95	-2.11
	DF_2	-0.51	-4.45	-2.60	-3.87	-2.82	-2.11
	DF_3	-2.15	-4.81	-3.97	-4.03	-4.22	-2.17
	DF_4	-3.21	-4.80	-5.30	-4.02	-5.20	-2.71
Lower Saxony	DF_0	1.06	-1.37	-0.75	-4.84	0.02	-3.54
	DF_1	-0.55	-2.60	-2.89	-4.83	-0.88	-3.75
	DF_2	-0.87	-3.80	-2.95	-4.83	-2.25	-3.77
	DF_3	-2.32	-4.01	-2.97	-4.81	-2.67	-4.00
	DF_4	-3.49	-4.01	-3.78	-4.81	-2.71	-4.48
Bremen	DF_0	1.48	-1.55	-2.13	-7.29	0.83	-3.81
	DF_1	-0.64	-3.43	-4.23	-7.31	-0.05	-4.05
	DF_2	-0.66	-4.53	-3.54	-7.67	-2.23	-4.58
	DF_3	-2.57	-4.84	-3.56	-7.26	-3.19	-4.69
	DF_4	-3.57	-4.85	-3.69	-6.79	-3.24	-4.72
Northrhine-Westfalia	DF_0	1.31	-1.57	-0.27	-4.94	0.65	-3.18
	DF_1	-0.37	-2.84	-4.35	-5.09	-0.67	-3.71
	DF_2	-0.82	-3.71	-4.11	-5.22	-1.86	-3.78
	DF_3	-2.43	-3.93	-3.75	-5.60	-2.67	-3.99
	DF_4	-3.50	-3.89	-4.20	-5.77	-2.72	-5.02
Hesse	DF_0	1.19	-1.61	-1.85	-3.94	-0.80	-2.13
	DF_1	-0.32	-2.72	-3.08	-3.95	-1.52	-2.22
	DF_2	-0.77	-3.61	-2.75	-4.14	-2.09	-2.05
	DF_3	-2.60	-3.76	-3.63	-4.35	-2.21	-2.62
	DF_4	-3.71	-3.73	-4.31	-4.36	-2.48	-3.82
Rhineland-Palatinate	DF_0	1.27	-1.72	-2.14	-6.27	0.30	-4.30
	DF_1	-0.26	-2.82	-2.61	-6.34	-0.58	-4.49
	DF_2	-0.87	-3.42	-2.61	-6.60	-1.85	-4.83
	DF_3	-2.73	-3.59	-5.08	-6.60	-3.84	-4.83
	DF_4	-3.57	-3.53	-5.08	-6.61	-4.11	-5.02
Saar	DF_0	1.82	-1.69	-0.90	-3.90	-0.11	-3.91
	DF_1	-0.42	-3.53	-3.09	-3.92	-1.39	-4.09
	DF_2	-0.95	-3.97	-2.97	-4.07	-3.60	-4.11
	DF_3	-2.19	-4.41	-3.33	-4.27	-3.60	-4.27
	DF_4	-2.96	-4.46	-4.73	-4.28	-4.04	-4.36
Baden-Württemberg	DF_0	1.59	-1.74	-1.43	-4.36	0.31	-3.37
	DF_1	-0.40	-3.49	-1.46	-4.49	-0.86	-3.64
	DF_2	-0.64	-4.57	-2.10	-4.53	-2.14	-3.64
	DF_3	-2.11	-4.78	-4.12	-4.70	-2.61	-3.98
	DF_4	-3.07	-4.75	-4.08	-5.27	-2.86	-4.70
Bavaria	DF_0	1.16	-1.49	0.12	-4.84	-0.91	-1.71
	DF_1	-0.23	-2.59	-1.02	-5.05	-1.60	-1.93
	DF_2	-1.05	-3.23	-3.45	-5.07	-2.42	-1.74
	DF_3	-2.59	-3.37	-3.43	-5.05	-3.21	-2.03
	DF_4	-3.59	-3.35	-3.94	-5.15	-3.27	-2.89
Berlin	DF_0	1.22	-1.50	-3.03	-2.98	-0.15	-4.62
	DF_1	-0.41	-2.86	-2.90	-3.02	-0.73	-4.68
	DF_2	-0.71	-4.49	-2.26	-3.57	-1.87	-5.33
	DF_3	-2.47	-4.59	-2.76	-4.02	-3.59	-5.37
	DF_4	-3.25	-4.69	-4.81	-4.02	-3.77	-5.37
West Germany	DF_0	1.28	-1.56	-	-	-	-
	DF_1	-0.33	-2.81	-	-	-	-
	DF_2	-0.81	-3.59	-	-	-	-
	DF_3	-2.43	-3.83	-	-	-	-
	DF_4	-3.44	-3.81	-	-	-	-

Notes: Test statistics DF_0, DF_1, DF_2, DF_3 and DF_4 as defined in the text, with (non-augmented) Dickey-Fuller test ($\gamma, \gamma' = 0$ in equations (13) and (15)) for first-order tests and augmented Dickey-Fuller test (unrestricted γ, γ' in equations (13) and (15)) for second-order tests. Critical values for rejecting the null hypothesis of a unit root at the 5 %- (10 %-) level are about -1.95 (-1.60) for DF_0, -3.00 (-2.63) for DF_1, and -3.60 (-3.24) for DF_2 (see Fuller 1976, p. 373, Table 8.5.2, for n=25). For DF_3 and DF_4 no such critical values are available, but it is certain that they will be higher in absolute terms than the ones for DF_2.

The results of the Engle/Granger-estimates for 1960-89 are shown in Table 5: for that subset of states which was also included in the larger sample, the results have the same characteristics as in Table 2 - thus indicating that no dramatic structural change need to be assumed between the decade of the 'German economic miracle' and later years. Note, in particular, that the standard error of the long-run estimates is again very small, on average even somewhat smaller than for the larger sample period. Hence our conclusion concerning the magnitude of regional wage differentiation receives additional support.

As to co-integration, the message of the two statistics DW and DF is again not conclusive. Most DW-statistics are higher in Table 5 than in Table 2, but so are the critical values at a larger range of indifference stretching from 0.747 to 1.156. The DF-statistics point towards rejecting the null hypothesis (i.e. a unit root in the residuals) in five out of eleven cases, if conventional significance levels (1 to 10 %) are applied, but another three states' estimates have DF-values at above 2.25 in absolute terms which would indicate significance at about the 20-30 %-level. The autoregression parameter ρ falls well below 0.8 in all estimates except the one for Berlin which is different in that it is the only state where the data speak somewhat conclusively in favour of a unit root and thus against co-integration (DW = 0.135, DF = 0.900): as Berlin is the only one out of the eleven West German states which has been geographically isolated over the sample period and which may thus have developed more of an 'autonomous' labour market, this result seems to have some economic significance. We shall return to this matter below.

As to the short-run dynamics, β_0 is again very close to one; the coefficient of the specifically regional labour market variable (γ_0) turns out 'correctly' signed in nine out of eleven cases, but in eight of them not significantly different from zero at the 5 %-level. Note that the only case in Table 2 of a γ_0 close to significance (Hesse) does not stand out any more when the sample is restricted to the time after the 1950s. As to the alternative specifications involving the absolute difference of unemployment rates ($\Delta(u^i-u)$) or the difference in output (employment) ratios ($\Delta(OR^i-OR)$, $\Delta(ER^i-ER)$), the diffuse picture of Table 2 is basically repeated with significant and 'correctly' signed parameters only in two cases for $\Delta(OR^i-OR)$ and in just one case for $\Delta(ER^i-ER)$.

Table 5 - Estimated Coefficients of the Engle/Granger-Model for West German States 1960-1989

	long-run (1st step)							short-run (2nd step)					short-run alternatively		
	const	log w	$\bar{R}^2$	SE(%)	DW	DF	ρ	Δlog w	$\Delta\log(u^i/u)$	v_{-1}	$\bar{R}^2$	DW	$\Delta(u^i-u)$	$\Delta(OR^i-OR)$	$\Delta(ER^i-ER)$
Schleswig-Holstein	-0.0400	1.0121	0.9987	2.27	0.7412	-2.4187	0.6296	1.0278	-0.0353	-0.4126	0.7964	1.8822	0.0017	0.1904	0.3582
	(0.0145)	(0.0067)						(0.0529)	(0.0479)	(0.1757)			(0.0179)	(0.2474)	(0.3176)
Hamburg	0.0805	1.0150	0.9993	1.67	0.6615	-2.0098	0.6879	1.0188	-0.0177	-0.3009	0.8629	2.0482	-0.0033	0.0166	-0.0507
	(0.0107)	(0.0049)						(0.0332)	(0.0265)	(0.1883)			(0.0062)	(0.2243)	(0.1467)
Lower Saxony	-0.0251	1.0163	0.9998	0.93	0.5487	-1.9585	0.7364	1.0098	0.0143	-0.2745	0.9536	1.7156	0.0068	-0.1835	-0.1597
	(0.0060)	(0.0028)						(0.0170)	(0.0215)	(0.1500)			(0.0067)	(0.1474)	(0.1478)
Bremen	-0.0069	1.0202	0.9991	1.88	0.5035	-2.2802	0.7164	0.9962	-0.0082	-0.2867	0.8479	1.3225	0.0000	-0.1391	0.1729
	(0.0121)	(0.0056)						(0.0322)	(0.0123)	(0.1453)			(0.0044)	(0.1731)	(0.0826)
Northrhine-Westfalia	0.0916	0.9739	0.9999	0.45	1.3495	-4.0416	0.2875	0.9764	-0.0076	-0.7915	0.9800	1.9665	0.0009	0.0538	-0.0600
	(0.0029)	(0.0013)						(0.0110)	(0.0139)	(0.2009)			(0.0054)	(0.1407)	(0.2084)
Hesse	-0.0200	1.0121	0.9998	0.79	0.3522	-2.8459	0.7185	1.0208	-0.0055	-0.3264	0.9814	1.7144	0.0007	0.0134	0.0051
	(0.0051)	(0.0023)						(0.0119)	(0.0120)	(0.1147)			(0.0068)	(0.1002)	(0.1272)
Rhineland-Palatinate	-0.1063	1.0330	0.9999	0.76	0.9431	-3.5386	0.4595	1.0385	0.0254	-0.6562	0.9714	1.8780	0.0115	0.1850	-0.1439
	(0.0049)	(0.0023)						(0.0151)	(0.0104)	(0.1560)			(0.0090)	(0.1374)	(0.1042)
Saar	0.0276	0.9958	0.9987	2.25	0.4410	-3.1591	0.6573	0.9740	-0.0180	-0.3523	0.8618	2.5637	0.0013	0.3099	0.0637
	(0.0144)	(0.0066)						(0.0329)	(0.0163)	(0.1219)			(0.0059)	(0.1560)	(0.1490)
Baden-Württemberg	-0.0960	1.0386	0.9999	0.60	1.5392	-4.5862	0.1871	1.0578	-0.0082	-0.8840	0.9757	2.1629	0.0076	0.1339	0.0383
	(0.0038)	(0.0018)						(0.0149)	(0.0067)	(0.1915)			(0.0046)	(0.1305)	(0.1328)
Bavaria	-0.1569	1.0312	0.9999	0.67	2.0375	-5.6075	-0.0527	1.0283	-0.0090	-1.1445	0.9538	1.9475	0.0028	0.1515	0.1429
	(0.0043)	(0.0020)						(0.0173)	(0.0182)	(0.2097)			(0.0087)	(0.2127)	(0.1766)
Berlin (West)	-0.0909	1.0333	0.9999	2.48	0.1351	-1.4835	0.9003	1.0304	-0.0180	-0.0855	0.9242	1.3350	-0.0011	-0.0622	0.0331
	(0.0159)	(0.0073)						(0.0247)	(0.0152)	(0.0777)			(0.0054)	(0.1440)	(0.1092)

Notes: $\bar{R}^2$ = adjusted R^2; SE = standard estimation error; DW = Durbin-Watson-statistic; DF = Dickey-Fuller-statistic for first-order autoregression of residuals; ρ = estimated coefficient of first-order autoregression of residuals; v_{-1} = lagged error term from first step of EG-model. Table based on estimates of equation (11); first step specified by equation (9) with θ_r = 0, second step by equation (10). Standard estimation errors of coefficients in parenthesis. For DW, the critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about 1.156 (range of indifference from 0.747 to 1.156). See Sargan, Bhargava (1983), p. 157, Table I for T=31 and n=1. For DF, critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about -3.365, at the 10 %-level about -3.066 and at the 15 %-level about -2.864; see Phillips, Ouliaris (1990), p. 190, Table II for n=1.

What makes the results of Table 5 interesting in their own right are the cases of the two additional (small) states Berlin and the Saar. Both states seem to stand out in that their estimates' standard errors are relatively large. Yet looking over all state results in Table 5, it becomes evident that all 'small' states, say, those with a population of less than 2.5 million (i.e. Schleswig-Holstein, Hamburg, Bremen, the Saar, and Berlin), have rather high estimation errors and most of them also rather low DW- and DF-statistics (in absolute terms). In fact, this makes one suspicious whether the explanatory power of the estimates and the co-integration properties may not be to some extent a mere reflection of the microcensus methodology used to calculate the wage data which, as described above,[41] may lead to an endogeneity bias depending on the size of the state. To make a rough-and-ready check whether such a bias ought to be a cause for concern, we carried out the EG-procedure for all possible combinations of the state data in the period 1960-89, without any recourse to the data for West Germany as a whole. Technically, this means that, in equation (11), the aggregate variables were step-by-step replaced by the respective variables for all other states except the one whose wage was to be explained. Table 6 presents the matrix of the standard estimation errors and the respective DW-statistics for all the 110 cross state estimates which cannot be affected by any endogeneity bias since they are based on disjunct samples of data. Comparing the (unweighted) averages of the 10 standard estimation errors and the respective DW-statistics of any state i, the familiar interstate pattern emerges, with the four small states having relatively high estimation errors and low DW. The same pattern can be discerned for any of the rows and columns of the matrix. All this gives one a fair amount of confidence that the endogeneity bias, if present at all, does not play a significant or even decisive role both with respect to the explanatory power of the estimates and to their co-integration properties.

Up to this point, we strictly stayed within the framework of the model as set out by equation (11), with the long-run effect of a regional labour market disequilibrium on the regional wage level assumed to be zero

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[41] See Part 3 above.

Table 6 - Standard Estimation Error and Durbin-Watson Statistic for 110 Cross State Estimates 1960-1989 (Engle/Granger-Model, First Step)

A. Standard Estimation Error (in %)	SH	HA	LS	BM	NW	HS	RP	SA	BW	BV	BL	Average	FRG
Schleswig-Holstein	.	1.79	2.06	3.10	2.38	2.21	2.37	3.64	2.31	2.42	2.40	2.47	2.27
Hamburg	1.80	.	1.39	2.49	1.68	1.40	1.67	3.45	1.76	1.77	1.92	1.93	1.67
Lower Saxony	2.06	1.75	.	1.75	1.39	1.22	1.23	1.79	0.89	0.74	1.22	1.40	0.93
Bremen	3.12	1.76	1.70	.	1.90	1.90	2.17	3.05	1.98	2.23	2.95	2.28	1.88
Northrhine-Westfalia	2.29	1.61	0.85	1.81	.	0.73	0.70	2.35	0.76	0.89	2.25	1.42	0.45
Hesse	2.21	1.40	0.73	1.89	0.76	.	0.93	3.00	0.74	0.97	1.87	1.45	0.79
Rhineland-Palatinate	2.41	1.70	1.24	2.20	0.75	0.95	.	2.55	1.05	0.87	2.58	1.63	2.04
Saar	3.58	3.38	2.99	2.98	2.41	2.96	2.46	.	2.60	2.47	4.44	3.03	2.10
Baden-Württemberg	2.37	1.80	1.04	2.02	0.81	0.76	1.06	2.71	.	0.96	2.29	1.58	0.60
Bavaria	2.47	1.80	1.20	2.25	0.94	0.99	0.87	2.56	0.95	.	2.59	1.66	0.67
Berlin	2.45	1.96	1.83	2.98	2.38	1.91	2.58	4.60	2.28	2.59	.	2.56	2.48
B. Durbin-Watson-Statistic													
Schleswig-Holstein	.	1.18	0.99	0.37	0.81	0.77	0.82	0.48	0.81	0.92	0.83	0.80	0.74
Hamburg	1.18	.	0.99	0.23	0.94	0.74	0.75	0.49	0.76	0.79	0.48	0.74	0.66
Lower Saxony	0.99	0.99	.	0.44	1.13	1.05	0.82	0.37	1.16	0.79	0.34	0.81	0.55
Bremen	0.37	0.23	0.44	.	0.68	0.45	0.49	0.51	0.58	0.61	0.25	0.46	0.50
Northrhine-Westfalia	0.81	0.94	1.13	0.68	.	1.04	1.36	0.27	1.67	2.07	0.20	1.02	1.35
Hesse	0.77	0.74	1.05	0.45	1.04	.	0.77	0.37	1.28	0.96	0.17	0.76	0.35
Rhineland-Palatinate	0.82	0.75	0.82	0.49	1.36	0.77	.	0.46	1.32	1.68	0.25	0.77	0.94
Saar	0.48	0.49	0.37	0.51	0.27	0.37	0.46	.	0.51	0.73	0.21	0.44	0.44
Baden-Württemberg	0.81	0.76	1.16	0.58	1.67	1.28	1.32	0.51	.	2.07	0.18	1.04	1.54
Bavaria	0.92	0.79	0.79	0.61	2.07	0.96	1.68	0.73	2.07	.	0.29	1.09	2.04
Berlin	0.83	0.48	0.34	0.25	0.20	0.17	0.25	0.21	0.18	0.29	.	0.32	0.14

Notes: SH = Schleswig-Hostein; HA = Hamburg; LS = Lower Saxony; BM = Bremen; NW = Northrhine-Westfalia; HS = Hesse; RP = Rhineland-Palatinate; SA = Saar; BW = Baden-Württemberg; BV = Bavaria; BL = Berlin; FRG = Federal Republic of Germany. Estimates corresponding to equation (9) with $\theta_r = 0$.

(θ_γ=0). As the tests for integration and co-integration have shown, the data are by and large compatible with this model, although some ambiguities remain. In particular, it is difficult to decide on basis of the test statistics (i) whether the proxies for the labour market disequilibrium (z^i-z) are integrated of degree one or stationary, and (ii) whether the wage level in any region i is co-integrated with the aggregate wage level or not. In the spirit of equation (11), we left question (i) open and assumed question (ii) to be answered in favour of co-integration. If we instead take question (i) to be answered in favour of integration of degree one and question (ii) against co-integration - both assumptions again not incompatible with the test statistics -, we can estimate equation (8) as a test for co-integration of the regional wage with a linear combination of the aggregate wage and the regional labour market disequilibrium. Note that the move from (11) to (8) concerns only the long run, i.e. θ_γ is not constrained to zero anymore; the specification of the short-run dynamics remains unchanged, and so do the assumptions with respect to the time series properties of the differenced variables ($\Delta\log(w^i)$, $\Delta\log(w)$ and $\Delta(z^i-z)$) which are all taken to be stationary, a realistic assumption as the second-order tests for integration have confirmed.

The Tables 7 and 8 present our estimates of the EG-model as specified by equation (8) for the periods 1950-89 and 1960-89 respectively; the format of the tables is analogous to that of the corresponding Tables 2 and 5.[42] Focusing on the differences between the results of the respective equations (11) and (8), five points are worth noting:

(i) As to the 'explanatory power' of the additional variable $\log(u^i/u)$ in the first step of the EG-procedure, the picture is mixed. For the period 1950-89, the long-run coefficient θ_γ has the expected negative sign in five out of nine cases, but only in two of them, it is also significant at the 5 %-level (Schleswig-Holstein, Bavaria); for the shorter period

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[42] The only major difference between the two sets of tables is that equation (8) has not been estimated with the difference in output (employment) ratios as a proxy for labour market disequilibrium, since they are stationary by definition.

Table 7 - Estimated Coefficients of the Engle/Granger-Model for West German States 1950-1989

	long-run (1st step)									short-run (2nd step)					
	const	log w	$\log(u^i/u)$	$\bar{R}^2$	SE(%)	DW	DF	ρ	(u^i-u)	$\Delta\log w$	$\Delta\log(u^i/u)$	v_{-1}	$\bar{R}^2$	DW	$\Delta(u^i-u)$
Schleswig-Holstein	0.0109	0.9953	-0.0802	0.9993	2.29	1.5310	-4.8434	0.2259	-0.0028	1.0165	-0.0310	-0.8467	0.7082	1.9707	0.0001
	(0.0290)	(0.0108)	(0.0322)						(0.0022)	(0.0521)	(0.0451)	(0.1768)			(0.0075)
Hamburg	0.0872	1.0117	-0.0079	0.9994	2.11	0.6090	-3.6021	0.5971	0.0005	0.9824	0.0296	-0.5384	0.8380	1.6494	0.0031
	(0.0078)	(0.0040)	(0.0098)						(0.0018)	(0.0292)	(0.0187)	(0.1261)			(0.0041)
Lower Saxony	-0.0417	1.0206	0.0307	0.9998	1.19	1.0145	-3.5704	0.4853	-0.0019	1.0084	0.0252	-0.5368	0.8840	2.0144	-0.0020
	(0.0085)	(0.0025)	(0.0229)						(0.0020)	(0.0232)	(0.0299)	(0.1721)			(0.0051)
Bremen	0.0508	0.9993	-0.0176	0.9993	2.35	0.3843	-1.8854	0.8147	0.0094	0.9761	-0.0062	-0.2335	0.8117	1.6772	0.0038
	(0.0131)	(0.0044)	(0.0164)						(0.0015)	(0.0302)	(0.0125)	(0.1095)			(0.0034)
Northrhine-Westfalia	0.0905	0.9714	0.0333	0.9998	1.05	1.0337	-5.2057	0.3423	0.0125	0.9987	0.0201	-0.6414	0.9430	1.6824	-0.0009
	(0.0057)	(0.0033)	(0.0077)						(0.0023)	(0.0180)	(0.0177)	(0.1626)			(0.0052)
Hesse	0.0019	1.0037	0.0192	0.9994	2.05	0.4060	-2.5534	0.7559	-0.0161	0.9952	-0.0336	-0.1646	0.8724	1.5706	-0.0258
	(0.0105)	(0.0040)	(0.0259)						(0.0047)	(0.0253)	(0.0223)	(0.1064)			(0.0100)
Rhineland-Palatinate	-0.1060	1.0316	-0.0136	0.9997	1.55	0.5827	-2.6850	0.6952	-0.0098	1.0313	-0.0164	-0.3211	0.8944	1.7909	-0.0044
	(0.0141)	(0.0053)	(0.0184)						(0.0027)	(0.0245)	(0.0132)	(0.1285)			(0.0061)
Baden-Württemberg	-0.0083	1.0098	0.0323	0.9995	2.03	0.6432	-3.6518	0.5890	-0.0123	0.9931	0.0016	-0.3835	0.8412	2.0767	-0.0115
	(0.0195)	(0.0058)	(0.0114)						(0.0019)	(0.0290)	(0.0139)	(0.1180)			(0.0067)
Bavaria	-0.1215	1.0098	-0.0606	0.9995	1.89	0.5727	-3.2283	0.6485	-0.0093	0.9844	-0.0347	-0.4029	0.8360	1.8000	-0.0106
	(0.0081)	(0.0060)	(0.0312)						(0.0055)	(0.0276)	(0.0283)	(0.1204)			(0.0127)

Notes: $\bar{R}^2$ = adjusted R^2; SE = standard estimation error; DW = Durbin-Watson-statistic; DF = Dickey-Fuller-statistic for first-order autocorrelation of residuals; ρ = estimated coefficient of first-order autoregression; v_{-1} = lagged error term from first step of EG-model. Table based on estimates of equation (11): first step specified by equation (9), second step by equation (10). Standard estimation errors of coefficients in parenthesis. For DW, the critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about 1.172 (range of indifference from 0.465 to 1.172); see Sargan, Bhargava (1983), p. 157, Table I for T=51 and n=3. For DF, critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about -3.768, at the 10 %-level about -3.449 and at the 15 %-level about -3.265; see Phillips, Ouliaris (1990), p. 190, Table II b for n=2.

Table 8 - Estimated Coefficients of the Engle/Granger-Model for West German States 1960-1989

	long-run (1st step)									short-run (2nd step)					
	const	log w	$\log(u^i/u)$	$\bar{R}^2$	SE(%)	DW	DF	ρ	(u^i-u)	$\Delta \log w$	$\Delta \log(u^i/u)$	v_{-1}	$\bar{R}^2$	DW	$\Delta(u^i-u)$
Schleswig-Holstein	0.0243	0.9902	-0.0905	0.9990	2.06	0.9421	-2.9214	0.5232	-0.0388	1.0178	-0.0666	-0.5142	0.8090	1.8718	-0.0126
	(0.0279)	(0.0103)	(0.0346)						(0.0126)	(0.0511)	(0.0469)	(0.1900)			(0.0180)
Hamburg	0.0575	1.0256	-0.0486	0.9997	1.11	1.4742	-3.9147	0.2271	-0.0084	1.0127	-0.0039	-1.0164	0.9216	1.6692	-0.0030
	(0.0081)	(0.0037)	(0.0081)						(0.0019)	(0.0251)	(0.0194)	(0.2082)			(0.0054)
Lower Saxony	-0.0378	1.0154	0.0718	0.9999	0.77	1.0256	-2.9195	0.4858	-0.0050	1.0098	0.0294	-0.4358	0.9566	1.6064	0.0058
	(0.0060)	(0.0023)	(0.0195)						(0.0040)	(0.0164)	(0.0217)	(0.1897)			(0.0067)
Bremen	-0.0270	1.0240	0.0206	0.9992	1.86	0.6830	-2.7479	0.6176	0.0069	0.9963	-0.0058	-0.2744	0.8431	1.3111	0.0011
	(0.0191)	(0.0062)	(0.0153)						(0.0021)	(0.0327)	(0.0132)	(0.1607)			(0.0043)
Northrhine-Westfalia	0.0901	0.9727	0.0186	0.9999	0.44	1.5572	-4.3410	0.1967	-0.0005	0.9776	0.0005	-0.8859	0.9819	1.9643	0.0006
	(0.0029)	(0.0015)	(0.0109)						(0.0020)	(0.0105)	(0.0131)	(0.1995)			(0.0054)
Hesse	-0.0088	1.0097	0.0297	0.9999	0.69	0.5661	-2.4537	0.6834	0.0079	1.0160	0.0034	-0.3532	0.9815	1.6121	0.0028
	(0.0057)	(0.0022)	(0.0096)						(0.0037)	(0.0120)	(0.0117)	(0.1304)			(0.0072)
Rhineland-Palatinate	-0.0921	1.0283	0.0255	0.9999	0.72	1.0514	-3.5768	0.4252	0.0090	1.0344	0.0342	-0.6712	0.9669	1.8001	0.0143
	(0.0079)	(0.0030)	(0.0116)						(0.0049)	(0.0164)	(0.0121)	(0.1965)			(0.0092)
Saar	0.0540	0.9997	-0.0700	0.9996	1.25	1.6981	-5.0159	0.1022	0.0111	0.9776	-0.0422	-0.8914	0.9064	2.1648	0.0034
	(0.0087)	(0.0037)	(0.0087)						(0.0073)	(0.0272)	(0.0132)	(0.1804)			(0.0063)
Baden-Württemberg	-0.1067	1.0418	-0.0047	0.9999	0.60	1.5441	-4.4880	0.1942	0.0029	1.0606	-0.0103	-0.8909	0.9759	2.1001	0.0100
	(0.0117)	(0.0038)	(0.0049)						(0.0020)	(0.0150)	(0.0067)	(0.1966)			(0.0047)
Bavaria	-0.1630	1.0359	0.0219	0.9999	0.65	2.3840	-6.5804	-0.2413	0.0029	1.0295	0.0073	-1.3169	0.9608	1.9539	0.0054
	(0.0055)	(0.0034)	(0.0131)						(0.0023)	(0.0160)	(0.0169)	(0.2091)			(0.0085)
Berlin (West)	-0.0307	1.0120	-0.0682	0.9996	1.26	0.7363	-2.6364	0.6136	-0.0234	1.0318	-0.0256	-0.3816	0.9383	1.3400	-0.0030
	(0.0105)	(0.0044)	(0.0076)						(0.0050)	(0.0216)	(0.0129)	(0.1332)			(0.0056)

Notes: $\bar{R}^2$ = adjusted R^2; SE = standard estimation error; DW = Durbin-Watson-statistic; DF = Dickey-Fuller-statistic for first-order autocorrelation of residuals; ρ = estimated coefficient of first-order autoregression; v_{-1} = lagged error term from first step of EG-model. Table based on estimates of equation (11): first step specified by equation (9), second step by equation (10). Standard estimation errors of coefficients in parenthesis. For DW, the critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about 1.736 (range of indifference from 0.699 to 1.736); see Sargan, Bhargava (1983), p. 157, Table I for T=31 and n=3. For DF, critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about -3.768, at the 10 %-level about -3.449 and at the 15 %-level about -3.265; see Phillips, Ouliaris (1990), p. 190, Table II b for n=2.

1960-89, a negative sign appears in five out of eleven cases, with four of the corresponding coefficients being significant at the 5 %-level (Schleswig-Holstein, Hamburg, the Saar, Berlin).

(ii) All in all, the explanatory power of the long-run estimates as measured by $\bar{R}^2$ and SE seems to be slightly better than in the respective equations of the Tables 2 and 5. Note, however, that in those cases where θ_γ has a 'perverse' sign and a t-ratio greater than one in absolute terms - three cases for 1950-89 and as many as six for 1960-89 -, the improvement owes much to the power of a variable which, on a priori grounds, should better be left out altogether. Qualitatively the same applies to the co-integration statistics: although most of them are better than in the corresponding equations of Tables 2 and 5, the improvement at least partly reflects the power of a variable with an implausibly signed coefficient.[43]

(iii) In the Tables 7 and 8, the most remarkable long-term estimates appear to be those for the four small states Berlin, Hamburg, the Saar, and Schleswig-Holstein in the shorter sample period 1960-89. In these cases, allowing for a non-zero long-term multiplier θ_γ leads to quite dramatic improvements in explanatory power and in co-integration properties as well as to the 'expected' negative sign of θ_γ. The case of Berlin is particularly striking: whereas its DF- and DW-statistics in Table 5 clearly speak against co-integration, they do not any more in Table 8 - thus indicating that the labour market in the geographically isolated territorium of West Berlin is in fact likely to have had its own long-run wage dynamics, which was not unaffected by regional disequilibria.

(iv) Qualitatively similar results for the long-run multipliers are obtained when (z^i-z) is proxied by (u^i-u) instead of $\log(u^i/u)$. For 1950-89, θ_γ

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43 In addition, the critical values for rejecting a unit root in the residuals are higher (in absolute terms) than in the respective equations of Tables 2 and 5 since more than one variable is included on the right-hand side of the co-integration equation.

turns out negative at the 5 %-significance level in four out of nine, for 1960-89 in just three out of eleven cases.

(v) As could be expected, the short-run dynamics estimated in the second step of the EG-procedure looks fairly invariant to whether the long-run multiplier θ_γ is restricted to zero or not. As in the Tables 2 and 5, the coefficient γ_0 turns out mostly insignificant and often 'wrongly' signed. Also, most other statistical characteristics of the relevant equations remain basically unchanged.

Again, we shall try to obtain a rough quantitative impression of the magnitude of the wage effects implied by our estimates. Following the methodology underlying Table 3, we assume that $u = u_{-1} = u_{-1}^i$ and $u^i = 2u_{-1}^i$, i.e. from an initial state of equal regional and aggregate unemployment rates - say 1 or 4 % -, the regional rate doubles. Table 9 presents the <u>long-run</u> effect of this exogenous shock on the regional wage for all states in the sample period 1950-89 and 1960-89, both for the two alternative specifications with $\log(u^i/u)$ and (u^i-u) as the relevant disequilibrium variable.

By and large, the results confirm our prior conjecture that the interregional structure of wages is fairly rigid in the long run; only in very few cases - Schleswig-Holstein for both samples, Bavaria for 1950-89 as well as Hamburg, the Saar and Berlin for 1960-89 - do the effects surpass 4 %; in all cases, they remain well below 10 %, if at all 'correctly' signed. The picture does not look very different for the semilog-linear specification: if the original aggregate and regional unemployment rates are 1 %, the doubling of the regional rate reduces the regional wage by less than 2 % in most cases, with Schleswig-Holstein and Berlin being the minor exceptions for the sample 1960-89. At a much higher unemployment rate of 5 % to start with, the effect becomes larger, but - again except for Schleswig-Holstein and Berlin - remains below the 10 %-level. On average, the long-run wage effect is just 1.6 %, a small number by any reasonable standard.

So much for the econometric evaluation of interregional wage rigidity with respect to regional labour market disequilibria. Despite some inherent shortcomings of the data and despite some ambiguities in the details of

Table 9 - Long-Run Effects on the Regional Wage (in %) of a Doubling in the Regional Labour Market Disequilibrium

	1950-1989			1960-1989		
	Log-linear	Semilog-linear Model		Log-linear	Semilog-linear Model	
$u_{-1} = u_{-1}^{i} = \ldots$	0-100 %	1 %	5 %	0-100 %	1 %	5 %
Schleswig-Holstein	-5.56	-0.28	-1.40	-6.27	-3.88	-19.40
Hamburg	-0.55	0.05	0.25	-3.37	-0.84	-4.20
Lower Saxony	2.13	-0.19	-0.95	4.98	-0.50	-2.50
Bremen	-1.22	0.94	4.70	1.43	0.69	3.45
Northrhine-Westfalia	2.31	1.25	6.35	1.29	-0.05	-0.25
Hesse	1.33	-1.61	-8.05	2.06	0.79	3.95
Rhineland-Palatinate	-0.94	-0.98	-4.90	1.77	0.90	4.50
Saar	-	-	-	-4.85	1.11	5.55
Baden-Württemberg	2.24	-1.23	-6.15	-0.33	0.29	1.45
Bavaria	-4.20	-0.93	-4.65	1.52	0.29	1.45
Berlin (West)	-	-	-	-4.73	-2.34	-11.70
Arithmetic Average	-0.50	-0.33	-1.66	-0.59	-0.32	-1.61

Notes: Long-run effects calculated on basis of the respective estimates in Tables 7 and 8.

the empirical picture, it is safe to say that the bulk of the evidence points to just one conclusion: for all that matters in practice, the regional wage structure is very rigid with respect to regional labour market disequilibria, both in the short- and in the long-run. Whatever story one may tell about the extent of regional wage differentiation in a descriptive sense, the flexibility of this very structure in the relevant economic sense of reacting to market disequilibria is modest at best.

5. Intersectoral and Interindustrial Wage Rigidity

We shall tackle the question of intersectoral/-industrial wage rigidity with basically the same theoretical and econometric apparatus as the one we used for the analysis of interregional wage rigidity. Again, the starting-point is a general dynamic framework of the type specified in equation (7), with all relevant regional variables redefined in terms of sectors or industries. Again, the general dynamic form can be rewritten in alternative versions leading to the Engle-Granger (EG)- and the Wickens-Breusch (WB)-model, with basically all technical comments made on these models - both to theory and to estimation - equally applying to regions and industries/sectors. As it was argued above, an intersectoral labour market disequilibrium should be interpreted as a short- and medium-run, not as a long-run phenomenon. Hence, a specification like equation (8) can only make sense for regions, but not for sectors or industries as the relevant long-term multiplier (θ_γ) is not constrained to zero. This is why we shall exclusively work with the specification given by equation (11) which imposes the constraint $\theta_\gamma = 0$. Given all these considerations and the limitations of the data, $(z^i - z)$ is taken to be the difference of the output (employment) ratios between industry/sector i and the aggregate, with the output (employment) ratio defined as the (log-)deviation of output or employment in the relevant industry or sector from a (possibly higher-order) deterministic trend.[44]

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44 Remember that $(z^i - z)$ is thus trend stationary by definition.

For the empirical analysis of interindustry and intersectoral wage rigidity, we alternatively use two independant data sources, namely the so-called industry statistics (1) and the national accounts (2). We shall separately discuss the evidence destilled from each of these sources.

5.1. Estimates Based on Industry Statistics

Quarterly and annual reports on industry are the only data source in West Germany which provides a consistent time series of wage and production indices from 1950 to the present. As to the structural disaggregation, the statistics follow the so-called SYPRO-framework which, roughly speaking, divides total industry (excluding construction) into mining and manufacturing, the latter consisting of four broad industry groups: basic materials, investment goods, consumption goods and food industries. These groups are again subdivided into a total number of 35 single manufacturing branches. For technical reasons,[45] we shall restrict our analysis to five sectors - mining and the four manufacturing industry groups - and to a selection of eight important manufacturing industries which together comprise 60-70 % of all manufacturing employees at any time of the sample period.

As to the wage variable, we made parallel use of two different wage indices which are of interest for our purposes: (i) the index of hourly conntractual minimum wages (w_T with T for 'Tarif') and (ii) the index of hourly earnings (w_{EF} with EF for 'Effektiv'), both of an average worker in West German manufacturing industry. (i) w_T is based on a representative sample of important collective agreements in the respective sector. It is defined as the pure time rate per hour of a worker with maximum seniority (the highest 'age bracket' and the highest 'local pay unit'). No supplementary benefits and premia are included. (ii) w_{EF} is based on a representative sample of actual wage payments in establishments with more than 10 employees. It is defined as the gross earnings per hour which include the contractual minimum wage as well as

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[45] In particular, quite a few industry branches are so small that part of the data needed is not published or otherwise accessible due to reasons of data privacy protection.

all supplementary benefits and premia which can be attributed to working time, such as overtime premia, performance-related rates etc.; excluded are all benefits which are paid on an annual or irregular basis outside the working time schedule (e.g. holiday money, gratifications, bonuses etc.). Both indices have a distinct meaning: with the usual statistical caveats, w_T can be taken as a measure of the contractual minimum wage, w_{EF} as a measure of the 'effective' wage. The difference of the growth rate of w_{EF} and w_T can be interpreted as a rough measure of a wage drift which reflects both the difference between the effective and the minimum wage per ordinary working hour, and the earnings effect of overtime premia which raise the effective wage per 'total' working hour (including overtime).[46]

Note that our wage variables w_{EF} and w_T avoid two of the conceptual problems which have plagued the analysis in the case of interregional wage differentiation. First, they obviously allow for a seperate estimation of contractual as compared to effective wage rigidity across industries; second, being Laspeyres-wage indices, they filter out that part of wage growth which is a mere reflection of intersectoral movements of labour as a consequence of structural change. On the other hand, they are still possibly subject to an endogeneity bias due to an intersectoral overlapping of the microcensus data base used for the calculation of the indices, just as was the case in our interregional analysis. As there, however, appropriate tests using strictly disjunct sectoral units indicated that the bias from this source is likely to be negligible.[47]

The output ratio for any sector and for manufacturing as a whole was calculated on basis of the standard index of net industrial production, the employment ratio on basis of the number of dependent status employees. Both ratios are defined as the residuals from a regression of

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[46] For details, see i.a. the glossary in the bi-annual publication Statistisches Bundesamt, Lange Reihen zur Wirtschaftsentwicklung, latest issue 1990.

[47] The design of these tests was analogous to that for regions. See above in Part 4.

the log of the industrial production index (the level of employment) on a constant and three trend terms (linear, quadratic and cubic).[48]

The pretesting for integration was limited to the two wage variables since the output and employment ratios are trend stationary by definition. Table 10 presents two Dickey-Fuller-test-statistics (DF_1, DF_4) from autoregressions for the minimum and for the effective wage (in logs); DF_1 and DF_4 are defined as in Part 4,[49] involving just a constant term in the case of DF_1 and a constant term plus linear, quadratic and cubic trend terms in the case of DF_4. The picture for DF_1 is quite unambiguous: I(1) cannot be rejected at any of the usual significance levels as most of the DF-statistics are in absolute terms below one, none above 1.33. In contrast, I(2) is rejected in most cases at least at the 10 %- and often at higher levels. As to DF_4, again, trend stationarity is a good approximation of the data generation process, both for levels and for differences. At any rate, the wages in most industry groups and single branches of manufacturing seem to have about the same time series properties as the wage for manufacturing as a whole. Note also that the difference between the respective test statistics for the minimum and the effective wage is very small. Summing up, the test statistics do clearly not preclude the application of an EG-model of the form (11) with the wage index for total manufacturing as a proxy for the relevant aggregate wage level.[50]

Table 11 and 12 present the results of the two-step EG-estimates of equation (11), separately for the minimum and the effective wage. The tables cover 13 cases each, 5 for the industry groups, 8 for the selected

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48 Note that the estimates turned out to be very robust with respect to the precise method of calculation of the output (employment) ratio. In particular, adding a fourth-order or leaving out the third-order trend term did hardly change the results.

49 See in particular equations (13) to (15).

50 As in Part 4 for regions, the test statistics do also point to the co-integration equation as being a test setting for deterministic, not just for stochastic co-integration.

Table 10 - Tests for Integration, Selected Industries 1950-1989

Variable (in logs): Order of Integr.: Statistics:	Minimum Wage 1st DF_1	Minimum Wage 1st DF_4	Minimum Wage 2nd DF_1	Minimum Wage 2nd DF_4	Effective Wage 1st DF_1	Effective Wage 1st DF_4	Effective Wage 2nd DF_1	Effective Wage 2nd DF_4
Mining	-0.34	-2.49	-3.25	-4.28	-0.22	-3.04	-3.97	-4.69
Manufacturing:								
- basic materials	-0.15	-3.16	-3.26	-4.73	0.11	-3.89	-3.79	-4.52
- investment goods	-0.15	-2.66	-4.23	-5.49	0.15	-3.87	-4.12	-5.02
- consumption goods	-0.23	-3.66	-2.90	-4.07	-0.08	-4.33	-3.09	-4.58
- food products	-0.45	-3.59	-2.35	-3.36	-0.18	-3.89	-2.73	-4.04
Selected branches:								
- iron and steel	0.00	-2.38	-4.59	-5.87	0.18	-3.44	-4.66	-4.84
- chemicals	-0.30	-3.42	-3.38	-5.12	0.05	-3.99	-3.97	-5.23
- mechanical engineering	-0.08	-2.58	-4.44	-5.70	0.16	-3.68	-4.27	-5.15
- motor vehicles	-0.25	-2.89	-3.98	-5.21	0.11	-3.82	-4.28	-5.24
- electrical engineering	-0.12	-2.67	-4.20	-5.40	0.20	-3.51	-4.24	-5.29
- shipbuilding	-0.12	-2.88	-4.46	-5.55	0.08	-3.67	-4.77	-5.79
- textiles	-0.41	-2.70	-3.42	-4.51	0.02	-3.24	-4.12	-5.65
- clothing	-0.37	-2.91	-2.97	-3.94	-0.20	-4.15	-3.51	-5.28
Manufacturing total	-0.18	-3.20	-3.22	-4.40	0.06	-4.11	-3.92	-4.72

Notes: Test statistics DF_1 and DF_4 as defined in the text, with (non-augmented) Dickey-Fuller test ($\gamma,\gamma' = 0$ in equations (13) and (15)) for first-order tests and augmented Dickey-Fuller test (unrestricted γ,γ' in equations (13) and (15)) for second-order tests. Critical value for rejecting the null hypothesis of a unit root at the 5 %- (10 %-) level is about -2.93 (-2.60) for DF_1 (see Fuller 1976, p. 373, Table 8.5.2., for n=50). For DF_4, no such critical value is available, but if it were, it would be higher in absolute terms than -3.50 (-3.18), the critical value for DF_2 as defined in the text.

Table 11 - Estimated Coefficients of the Engle/Granger-Model for Selected Industries, Hourly Minimum Wages 1950-1989

	long-run (1st step)							short-run (2nd step)					
Industry Group	const	log w	$\bar{R}^2$	SE(%)	DW	DF	ρ	Δlog w	$\Delta(OR^i-OR)$	v_{-1}	$\bar{R}^2$	DW	$\Delta(ER^i-ER)$
Mining	0.2564	0.9625	0.9954	5.15	0.1386	-1.3374	0.9211	0.9459	0.0574	-0.0919	0.6503	1.6919	0.0961
	(0.0488)	(0.0105)						(0.0485)	(0.0846)	(0.0659)			(0.0712)
Basic materials	-0.1179	1.0200	0.9999	0.82	0.5849	-2.2494	0.7254	1.0244	0.0289	-0.2849	0.9490	2.3396	-0.0066
	(0.0078)	(0.0017)						(0.0157)	(0.0492)	(0.1468)			(0.0989)
Investment goods	0.1095	0.9788	0.9999	0.91	0.5186	-2.2436	0.7465	0.9836	-0.0151	-0.2660	0.9421	2.6229	-0.1485
	(0.0086)	(0.0018)						(0.0160)	(0.0518)	(0.1275)			(0.1000)
Consumption goods	-0.0554	1.0140	0.9998	1.20	0.4919	-2.3000	0.7545	1.0052	-0.0008	-0.2640	0.9001	2.1650	0.1251
	(0.0114)	(0.0024)						(0.0200)	(0.0590)	(0.1187)			(0.0845)
Food, beverages etc.	-0.1474	1.0354	0.9993	2.22	0.2654	-1.8514	0.8514	1.0187	0.1079	-0.1384	0.8392	1.8922	0.2947
	(0.0210)	(0.0045)						(0.0263)	(0.0546)	(0.0857)			(0.0745)
Selected Branches													
Iron and steel	0.0684	0.9806	0.9992	2.13	0.3750	-1.9617	0.8139	0.9854	0.0417	-0.1874	0.7968	1.7826	0.0284
	(0.0202)	(0.0043)						(0.0307)	(0.0309)	(0.1018)			(0.0954)
Chemicals	-0.4368	1.0790	0.9988	2.96	0.3057	-1.4014	0.8725	1.0812	-0.0412	-0.1414	0.7413	2.1622	0.0078
	(0.0281)	(0.0060)						(0.0435)	(0.0828)	(0.1075)			(0.0174)
Mechanical engineering	0.1245	0.9744	0.9999	0.90	0.5739	-1.9773	0.7490	0.9788	-0.0015	-0.2643	0.9320	2.5179	-0.0607
	(0.0085)	(0.0018)						(0.0170)	(0.0334)	(0.1427)			(0.0587)
Motor vehicles	0.0720	0.9920	0.9996	1.50	0.4421	-2.5793	0.7434	0.9959	0.0110	-0.2861	0.8911	2.2061	-0.0322
	(0.0142)	(0.0031)						(0.0249)	(0.0309)	(0.1172)			(0.0603)
Electrical engineering	0.1196	0.9765	0.9999	0.91	0.4510	-1.6467	0.8119	0.9790	-0.0404	-0.1956	0.9451	2.2061	-0.0615
	(0.0086)	(0.0018)						(0.0151)	(0.0359)	(0.1268)			(0.0472)
Shipbuilding	0.2060	0.9533	0.9998	1.04	0.6383	-2.3542	0.6914	0.9514	0.0137	-0.3178	0.8979	2.1204	0.0217
	(0.0099)	(0.0021)						(0.0197)	(0.0137)	(0.1441)			(0.0213)
Textiles	-0.0814	1.0188	0.9990	2.53	0.4734	-2.2442	0.7627	1.0095	-0.0409	-0.2579	0.6862	1.9053	-0.0794
	(0.0239)	(0.0051)						(0.0418)	(0.0845)	(0.1175)			(0.1327)
Clothing	0.0464+	1.0026	0.9988	2.76	0.2917	-2.3478	0.8078	0.9759	-0.0206	-0.2221	0.7738	1.5360	0.0316
	(0.0261)	(0.0056)						(0.0358)	(0.0572)	(0.0915)			(0.0861)

Notes: $\bar{R}^2$ = adjusted R^2; SE = standard estimation error; DW = Durbin-Watson-statistic; DF = Dickey-Fuller-statistic for first-order autocorrelation of residuals; ρ = estimated coefficient of first-order autoregression; v_{-1} = lagged error term from first step of EG-model. Table based on estimates of equation (11): first step specified by equation (9) with $\theta_r = 0$, second step by equation (10). Standard estimation errors of coefficients in parenthesis. For DW, the critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about 0.755 (range of indifference from 0.484 to 0.755); see Sargan, Bhargava (1983), p. 157, Table I for T=51 and n=1. For DF, critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about -3.365, at the 10 %-level about -3.066 and at the 15 %-level about -2.864; see Phillips, Ouliaris (1990), p. 190, Table II b for n=1.

Table 12 - Estimated Coefficients of the Engle/Granger-Model for Selected Industries, Effective Hourly Wages 1950-1989

	long-run (1st step)							short-run (2nd step)					
Industry Group	const	log w	$\bar{R}^2$	SE(%)	DW	DF	ρ	Δ log w	$\Delta(OR^i-OR)$	v_{-1}	$\bar{R}^2$	DW	$\Delta(ER^i-ER)$
Mining	0.1632	0.9719	0.9973	4.25	0.2052	-1.3779	0.9008	0.9719	0.2234	-0.1128	0.7985	1.5397	0.1948
	(0.0372)	(0.0081)						(0.0391)	(0.0767)	(0.0723)			(0.0668)
Basic materials	0.0407	0.9890	0.9999	0.81	1.1094	-4.7199	0.3562	0.9908	0.0336	-0.6818	0.9445	2.0007	0.0447
	(0.0071)	(0.0015)						(0.0156)	(0.0565)	(0.1584)			(0.1122)
Investment goods	-0.0297	1.0069	0.9999	0.88	1.0865	-3.7089	0.4524	0.9920	0.0684	-0.6190	0.9330	2.1994	-0.1547
	(0.0077)	(0.0017)						(0.0177)	(0.0652)	(0.1615)			(0.1299)
Consumption goods	-0.0432	1.0070	0.9995	1.83	0.3166	-1.6912	0.8473	0.9751	-0.0103	-0.1454	0.8802	1.7122	0.1239
	(0.0160)	(0.0035)						(0.0228)	(0.0747)	(0.1026)			(0.1074)
Food, beverages etc.	-0.1390	1.0327	0.9995	1.85	0.4860	-1.8374	0.7889	0.9962	0.0914	-0.1722	0.8400	1.7660	0.2364
	(0.0162)	(0.0035)						(0.0273)	(0.0637)	(0.1254)			(0.0936)
Selected Branches													
Iron and steel	0.2605	0.9418	0.9992	2.18	0.4876	-3.4448	0.6518	0.9554	0.0786	-0.3013	0.8397	1.5640	0.0475
	(0.0191)	(0.0042)						(0.0282)	(0.0333)	(0.1096)			(0.1012)
Chemicals	-0.1424	1.0258	0.9997	1.49	0.6046	-2.4247	0.7050	1.0194	0.0360	-0.2952	0.8820	2.0525	0.0089
	(0.0130)	(0.0028)						(0.0250)	(0.0563)	(0.1405)			(0.0114)
Mechanical engineering	-0.1523	1.0027	0.9998	1.32	0.6149	-2.6079	0.6920	0.9872	0.0847	-0.3509	0.9078	2.2837	0.0209
	(0.0115)	(0.0025)						(0.0214)	(0.0477)	(0.1264)			(0.0880)
Motor vehicles	-0.0838	1.0206	0.9996	1.67	0.4487	-2.3150	0.7618	0.9994	0.0239	-0.2654	0.8871	2.3201	-0.0806
	(0.0146)	(0.0032)						(0.0237)	(0.0331)	(0.1137)			(0.0633)
Electrical engineering	-0.0181	1.0070	0.9997	1.49	0.4207	-2.3830	0.7584	0.9836	-0.0019	-0.2744	0.9078	2.1276	-0.0713
	(0.0131)	(0.0028)						(0.0206)	(0.0056)	(0.1121)			(0.0729)
Shipbuilding	0.0219	0.9973	0.9997	1.35	0.9240	-3.3225	0.5373	0.9688	0.0423	-0.5814	0.8861	2.2059	0.0596
	(0.0118)	(0.0026)						(0.0245)	(0.0193)	(0.1468)			(0.0308)
Textiles	-0.0043	0.9980	0.9997	1.35	0.9527	-3.5515	0.5078	0.9830	0.0155	-0.5160	0.8428	2.0433	0.1189
	(0.0118)	(0.0526)						(0.0264)	(0.0591)	(0.1557)			(0.0910)
Clothing	-0.0497	1.0087	0.9990	2.64	0.3332	-1.5856	0.8503	0.9797	0.0405	-0.1405	0.7738	1.8903	0.0649
	(0.0232)	(0.0050)						(0.0339)	(0.0610)	(0.1057)			(0.0918)

Notes: $\bar{R}^2$ = adjusted R^2; SE = standard estimation error; DW = Durbin-Watson-statistic; DF = Dickey-Fuller-statistic for first-order autocorrelation of residuals; ρ = estimated coefficient of first-order autoregression; v_{-1} = lagged error term from first step of EG-model. Table based on estimates of equation (11): first step specified by equation (9) with $\theta_r = 0$, second step by equation (10). Standard estimation errors of coefficients in parenthesis. For DW, the critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about 0.755 (range of indifference from 0.484 to 0.755); see Sargan, Bhargava (1983), p. 157, Table I for T=51 and n=1. For DF, critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about -3.365, at the 10 %-level about -3.066 and at the 15 %-level about -2.864; see Phillips, Ouliaris (1990), p. 190, Table II b for n=1.

wage index for total manufacturing. The statistical message which can be derived from the table, may be condensed to three major points.

(i) As in the estimates for regions, the $\bar{R}^2$ is high all throughout - mostly above 0.999 - and the standard estimation error (SE) very small - usually well below 3 %, in some cases even below 1 %. The only notable exception is the mining industry, one of the major losers of long-run structural change, for which SE turns out to be much higher - 5.15 % for w_T and 4.25 % for w_E. The long-run multipliers θ_β fall again in a rather narrow interval around one [0.94; 1.08], which, however, is not quite as narrow as was the case in the regional part of our analysis. Again, the standard estimation errors of θ_β are extremely small so that, in the vast majority of cases, θ_β turns out to be significantly different from one at the usual confidence levels.

(ii) With respect to co-integration, the relevant test statistics DW and DF are - as usual - not conclusive. Nevertheless, a slight but systematic disparity between the results in the two tables is discernible: for the estimates with the minimum wage (Table 11), the DF-statistics do not allow in any case to reject the null hypothesis of no co-integration at any of the standard significance levels; on the other hand, the estimates with the effective wage (Table 12) lead in general to perceptibly higher DF-statistics and in five cases (out of 13), the null hypothesis can in fact be rejected at a 15 %- (or a higher) significance level. The DW-statistic qualitatively confirms this picture, although - again as usual - its discriminatory power is estremely low. A glance over the estimated autoregressive parameters ρ of the residuals confirms that no strong conclusions should be drawn: only in a few industries - most notably mining - is ρ so close to one that assuming anything else than a unit root would be far-fetched. For most other industries, ρ is somewhere between 0.7 and 0.85 for the minimum wage and between 0.5 and 0.8 for the effective wage - thus indicating that the tests allow for a wide range of autoregressive structures including a unit root. Only the slight disparity between the tables appears to be something of an established fact, with the data speaking a bit more against co-integration for w_T than for w_E. Apparently, sectoral minimum wages as set in collective agreements have a stronger element of what might be called 'autonomous inertia' which makes them less responsive to the economy-wide (deterministic or stochastic) trends than effective wages. Why this should be so is a matter of speculation not to be dealt with at this point. Note, however, that this 'autonomous inertia' does not imply anything like a genuine flexibility of collective bargaining with respect to specifically sectoral conditions; as we argued above, these conditions are necessarily of a transitory nature and thus cannot be the driving force behind a non-explained stochastic trend of the sectoral wage.

(iii) As to the short-run dynamics, the statistical picture has features which are quite familiar from the regional part of our analysis. The short-run coefficient β_0 of $\Delta\log(w)$ is again highly significant and close to one, usually in fact not far off the long-run coefficient θ_β as estimated in the respective first step. As usual, the coefficient $-(1-\alpha_1)$ of the lagged level term v_{-1} is of a similar magnitude as the autoregressive parameter ρ in the DF-test-estimate: for w_T, α_1 falls in the interval [0.7; 0.9], for w_E in the interval [0.5; 0.8].

Most important for evaluating the extent of intersectoral wage rigidity is, of course, the coefficient γ_0 of the disequilibrium variable $\Delta(OR^i-OR)$: for w_T, it turns out insignificant at the 5 %-level in all estimates and 'wrongly' signed in seven out of 13 cases. Replacing $\Delta(OR^i-OR)$ by $\Delta(ER^i-ER)$, the results are no better: again, β_0 is insignificant all throughout and implausibly signed in six out 13 cases. Hence there is not the slightest evidence of any reaction of sectoral minimum wages to transitory sectoral disequilibria. The picture looks a bit different for the estimates with w_E: using $\Delta(OR^i-OR)$ as disequilibrium variable, γ_0 has the expected sign in as many as eleven out of 13 cases, in five of which it is also significant at the 5 %-level; replacing $\Delta(OR^i-OR)$ by $\Delta(ER^i-ER)$, ten out of 13 estimates of γ_0 have the expected sign, albeit only three at the 5 %-significance level. Hence, although the evidence is far from conclusive at this point, the data seem to speak for a somewhat greater flexibility of effective sectoral wages to sectoral disequilibria than of minimum wages as set in collective bargaining.

To obtain again a quantiative impression of the extent of wage flexibility involved, Table 13 presents the impulse effect of a 10 %-drop of the sectoral output ratio relative to the aggregate one. Note that, by all standards, such a 10 %-drop is an altogether dramatic sectoral event

Table 13 - Impulse Effect on the Sectoral Wage (in %) of a 10 %-Decline in the Sectoral Output Ratio

	Minimum Wage	Effective Wage
Mining	-0.57	-2.23
Manufacturing	-0.66	-1.02
- basic material	-0.29	-0.34
- investment goods	0.15	-0.68
- consumption goods	0.00	0.10
- food products	-1.08	-0.91
Selected manufacturing		
- iron and steel	-0.42	-0.79
- chemicals	0.41	-0.36
- mechanical engineering	0.02	-0.85
- motor vehicles	-0.11	-0.24
- electrical engineering	0.40	0.02
- shipbuilding	-0.14	-0.42
- textiles	0.41	-0.16
- clothing	0.21	-0.41

Notes: Impulse effects calculated on basis of the respective estimates in Tables 11 and 12.

which goes beyond even the dimension of the most powerful business downturns of, say, 1967, 1974/75 and 1981/82. Two points can be distilled from the numbers in the table. Firstly, as discussed above, the flexibility of the minimum wage - if at all in the 'right' direction - is weaker than that of the effective wage. Secondly, none of the numbers indicates anything like a substantial impulse effect; after all, even the maximum of a 2.2 %-differential in wage growth temporarily emerging between mining and manufacturing as a reaction to a 10 %-drop of the respective output ratio is very small by any sensible standard, not to speak of the less than 1 %-reactions of wage growth in other sectors in analogous situations. Hence one can hardly escape the conclusion that the industrial wage structure is very rigid with respect to sectoral crises, probably even more so than the interregional wage structure with respect to regional labour market disequilibria.

5.2. Estimates Based on National Accounts Statistics

From 1960 to the present, the West German national accounts statistics provide a consistent data base of gross incomes from wages and salaries, of value added at constant prices ('output') and of the level of employment for broad sectors of economic activity as well as narrowly defined branches of manufacturing. Combined with sectorally disaggregated data on average annual working hours from a different, but compatibel source,[51] the national accounts statistics allow to calculate the average hourly gross remuneration of employees (including social security contributions of employers) which - for all that matters - is the broadest available measure of an effective wage.[52]

Of course, the national accounts statistics have some inherent shortcomings; in particular, the data are not based on an index calculation and thus reflect all effects of sectoral structural change on hourly earnings. Nevertheless, as a second independent data source for the empirical analysis of interindustrial and intersectoral wage rigidity, the national accounts have also quite distinctive advantages. First, they cover not only industry in the narrow sense (i.e. manufacturing and mining), but also agriculture, construction and the private service sectors. Second, given the extremely broad definition of the wage, they may yield estimates for something like an upper bound of flexibility to sectoral disequlibria; this is so to the extent that elements of remuneration other than actual wage payments serve as hidden instruments of flexibility. Finally, the national accounts data allow to define sectoral and economy-wide wages in a way so as to make them based on completely disjunct samples; this can be achieved by simply calculating the aggregate wage as the economy-wide average excluding the sector or industry branch under consideration.

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[51] The data on working hours has been provided by the Bundesanstalt für Arbeit. It is part of a data bank at the Institute of World Economics. Thanks are due to H. Klodt for guidance on this matter.

[52] Note that using a somewhat less broad measure (e.g., excluding social securitiy contributions) did hardly affect the results.

Keeping these data characteristics in mind, we use exactly the same model for the national accounts as for the industry index statistics, with the wage variable now defined as the average hourly gross earnings (including social security contributions of employers) of a dependent status employee and the output (employment) ratio defined as the residual from a regression of the log of value added at constant prices (the level of employment) on a constant and three trend terms (linear, quadratic and cubic).[53] Table 14 presents the results of the familiar EG-model (equation 11) for seven broad sectors of economic activity (agriculture and forestry, energy and mining, manufacturing, construction, trade and commerce, transport and communications, banking and insurance), with the relevant aggregate variables calculated for the whole privat for-profit sector of the economy excluding the sector under consideration. In turn, Table 15 covers 31 different branches of manufacturing, with the aggregate defined as total manufacturing excluding the branch under consideration. Note that the standard pre-testing of the wage variables for their degree of integration brought basically the same results as for the industry statistics, namely I(1) to be accepted and I(2) to be rejected for the vast majority of cases, including the respective aggregate wage variables. To save on space, we have not reprinted the test results in extra tables.[54]

Glancing over the results of the two tables and picking out the most characteristic features, which distinguish the estimates from those based on industry index statistics, the following three points are worth noting:

(i) By and large, the explanatory power of the long-run estimates is again very high, although not quite as high as before: in most cases, the $\bar{R}^2$ is below 0.999 though still above 0.990 whereas the SE falls in the range 1 to 6 %, on average being around 2.6 %. These numbers indicate that, in fact, the somewhat broader wage definition leaves more

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[53] Again, the precise definition of the trend did not matter much for the outcome of the estimates.

[54] Again, the data generation process could also be described by a higher-order deterministic trend, with the cubic term usually being decisive.

Table 14 - Estimated Coefficients of the Engle/Granger-Model for Seven Sectors, Hourly Earnings, West Germany 1960-1989

	long-run (1st step)							short-run (2nd step)					
	const	log w	$\bar{R}^2$	SE(%)	DW	DF	ρ	Δlog w	$\Delta(OR^i-OR)$	v_{-1}	$\bar{R}^2$	DW	$\Delta(ER^i-ER)$
Agriculture, forestry	-0.3452	0.9690	0.9943	5.12	0.1775	-1.0510	0.9142	0.9209	0.0777**	-0.1022**	0.6113	1.7119	-0.1327**
	(0.0344)	(0.0139)						(0.0504)	(0.0618)	(0.0903)			(0.2370)
Energy, mining	0.2444	1.0579	0.9971	3.95	0.2408	-2.0998	0.8148	1.0252	0.0635**	-0.2049*	0.8080	1.5254	-0.3588**
	(0.0264)	(0.0107)						(0.0433)	(0.1107)	(0.0976)			(0.2095)
Manufacturing	-0.0569	1.0853	0.9984	2.93	0.2448	-1.5083	0.8592	1.0397	0.1474**	-0.8792	0.8792	1.3011	-0.0154**
	(0.0201)	(0.0083)						(0.0326)	(0.1114)	(0.0985)			(0.1693)
Construction	0.1588	0.9344	0.9963	4.01	0.7588	-2.9395	0.5582	1.0198	0.5478	-0.4305	0.8067	2.1597	0.6267
	(0.0268)	(0.0108)						(0.0588)	(0.1634)	(0.1355)			(0.2440)
Trade, commerce	-0.2742	1.0371	0.9988	2.46	0.5345	-2.1010	0.7249	1.0237	-0.0026**	-0.3005	0.7401	1.2949	0.1286**
	(0.0167)	(0.0067)						(0.0414)	(0.2836)	(0.1529)			(0.3484)
Transport, communications	0.2746	0.9396	0.9975	3.29	0.2640	-1.2117	0.8800	0.9568	-0.1072**	-0.1442	0.8026	1.9077	-0.0785**
	(0.0219)	(0.0089)						(0.0411)	(0.2480)	(0.1152)			(0.2537)
Banking, insurance	0.3390	1.0004	0.9984	2.76	0.4654	-3.3949	0.6125	0.9546	0.2245*	-0.4300	0.8292	1.4874	0.5129
	(0.0185)	(0.0075)						(0.0340)	(0.1039)	(0.1135)			(0.1952)

Notes: $\bar{R}^2$ = adjusted R^2; SE = standard estimation error; DW = Durbin-Watson-statistic; DF = Dickey-Fuller-statistic for first-order autocorrelation of residuals; ρ = estimated coefficient of first-order autoregression; v_{-1} = lagged error term from first step of EG-model. Table based on estimates of equation (11): first step specified by equation (9) with θ_γ = 0, second step by equation (10). Standard estimation errors of coefficients in parenthesis. For DW, the critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about 1.156 (range of indifference from 0.747 to 1.156); see Sargan, Bhargava (1983), p. 152, Table I for T=31 and n=1. For DF, critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about -3.365, at the 10 %-level about -3.066 and at the 15 %-level about -2.864; see Phillips, Ouliaris (1990), p. 190, Table II b for n=1. Significance level of estimated coefficients (other than ρ): '**' not significantly different from zero at the 5 %- (or higher) level; '*' significantly different from zero at the 5 %-level, but not at the 1 %-level; all other coefficients significantly different from zero at the 1 %- (or higher) level.

Table 15 - Estimated Coefficients of the Engle/Granger-Model for 31 Manufacturing Industries, Hourly Earnings, West Germany 1960-1989

	long-run (1st step)							short-run (2nd step)					
	const	log w	$\bar{R}^2$	SE(%)	DW	DF	ϱ	Δlog w	$\Delta(OR^i-OR)$	v_{-1}	$\bar{R}^2$	DW	$\Delta(ER^i-ER)$
Chemicals	0.1888	1.0434	0.9995	1.65	1.6464	-4.3455	0.1766	1.0515	0.0642**	-0.8310	0.8204	1.9386	-0.2824**
Refineries	0.3628	1.0295	0.9955	5.03	0.4576	-2.4402	0.7104	1.0889	-0.0341**	-0.3259	0.5726	1.9035	-0.1517**
Plastic products	-0.1146	0.9937	0.9988	2.47	0.4996	-2.1646	0.7266	1.0034	0.0760**	-0.3279	0.8071	1.8025	-0.1352**
Rubber	0.1805	0.9463	0.9995	1.53	0.6147	-2.0016	0.7026	0.9560	-0.0829**	-0.2875	0.9128	1.6683	-0.1243*
Stone and minerals	-0.0159**	0.9982	0.9994	1.84	0.5803	-2.1153	0.7115	0.9945	0.0840**	-0.2698	0.8606	1.8361	0.2216
Ceramics	-0.0498*	0.9605	0.9980	3.12	0.3202	-1.9399	0.8001	1.0004	-0.1327**	-0.2301	0.8531	1.5034	-0.4085
Glass	-0.0338*	1.0037	0.9992	2.12	0.6500	-2.5650	0.6420	1.0148	0.0855**	-0.4101	0.8041	2.1630	0.2749*
Iron and steel	0.2682	0.9802	0.9977	3.41	1.0195	-3.7441	0.4089	0.9669	-0.0507**	-0.6549	0.6739	2.1626	0.2169**
Non-ferrous metals	0.0792	1.0045	0.9993	1.88	0.6410	-2.3781	0.6665	1.0045	0.0067**	-0.3456	0.8650	1.9235	-0.1574**
Foundries	0.1361	0.9722	0.9996	1.34	0.9854	-3.6699	0.4203	0.9652	0.0275**	-0.5971	0.9039	1.9961	0.0358**
Rolling mills	0.0984	0.9116	0.9997	1.14	1.4503	-3.9074	0.2707	0.9167	-0.0032**	-0.7657	0.9019	1.9266	0.1540**
Metal manufacturing	0.1183	0.9312	0.9980	3.03	0.4373	-1.8555	0.7723	0.9697	-0.0113**	-0.2757	0.8036	1.2901	0.0397**
Mechanical engineering	0.0640	1.0022	0.9997	1.19	0.7775	-2.2173	0.6247	1.0077	-0.0079**	-0.4159	0.9213	2.3090	0.0628**
Business machines	-0.1338	1.2076	0.9977	4.23	0.8496	-2.0815	0.6006	1.2551	0.0632**	-0.4636	0.6388	1.8773	0.0618**
Motor vehicles	0.0058**	1.0531	0.9993	2.01	1.1489	-3.8720	0.3551	1.0166	-0.0876**	-0.6905	0.8046	2.1891	0.0003**
Shipbuilding	0.0283*	1.0242	0.9989	2.44	1.0639	-3.2054	0.4565	1.0608	-0.0122**	-0.6015	0.7651	2.1438	-0.0176**
Air- and spacecraft	-0.0985	1.1402	0.9976	4.06	0.4707	-1.8779	0.7662	1.1600	-0.0516**	-0.2493	0.6424	1.8031	-0.0894**
Electrical engineering	-0.0517	1.0505	0.9997	1.28	2.0133	-5.2506	-0.0141	1.0559	-0.1428**	-1.0529	0.8900	1.8629	-0.1796**
Optical instruments	-0.0127**	0.9566	0.9994	1.66	0.8186	-2.8450	0.5655	0.9513	0.0171**	-0.4659	0.8228	2.1501	-0.0661**
Iron, tin and metal goods	-0.0861	1.0116	0.9993	2.01	0.3338	-1.5154	0.8370	1.0006	0.0782**	-0.2088	0.8869	1.4136	0.0207**
Musical instruments and toys	-0.3063	1.0283	0.9985	2.88	0.7362	-2.3402	0.6359	0.9852	0.1115**	-0.3313	0.6241	2.2142	-0.2293**
Wood working	-0.1741	1.0105	0.9991	2.21	0.5887	-2.3403	0.6848	1.0009	-0.0105**	-0.3314	0.7632	1.5738	-0.1777**
Wood processing	0.0665	0.8765	0.9983	2.67	0.4971	-2.3455	0.7018	0.8938	-0.0842**	-0.3706	0.7469	1.8586	0.1427**
Cellulose, pulp and paper	-0.0101	1.0305	0.9987	2.73	0.7095	-2.5969	0.6234	1.0412	-0.1026**	-0.4095	0.8065	1.9395	-0.0719**
Pulp and paper processing	-0.1934	1.0107	0.9996	1.55	0.3772	-2.5137	0.7325	1.0330	-0.0161**	-0.3088	0.9379	1.9070	0.0911**
Printing	0.1110	0.9313	0.9969	3.79	0.1833	-1.4286	0.8841	0.9402	-0.1178**	-0.1136**	0.7959	1.5280	-0.2232*
Leather	-0.0940	0.8841	0.9952	4.45	0.2493	-0.8833	0.9132	0.8519	0.0489**	-0.0996**	0.4849	1.9895	0.0540**
Textiles	-0.1636	0.9797	0.9998	1.04	1.2111	-3.5479	0.3770	0.9839	-0.1905	-0.4715	0.9363	2.4826	-0.3175
Clothing	-0.2133	0.9259	0.9997	1.18	1.0771	-4.6326	0.3093	0.9342	-0.0604**	-0.8167	0.9213	1.8494	0.1764
Food and beverages	0.2171	0.8080	0.9995	1.31	0.6900	-2.9665	0.5700	0.8096	-0.0054**	-0.4760	0.8813	1.9003	-0.0489**
Tobacco	-0.5186	1.2172	0.9988	3.07	1.1919	-3.4249	0.3948	1.2090	-0.1467**	-0.7278	0.5749	1.7137	-0.2378**

Notes: $\bar{R}^2$ = adjusted R^2; SE = standard estimation error; DW = Durbin-Watson-statistic; DF = Dickey-Fuller-statistic for first-order autocorrelation of residuals; ϱ = estimated coefficient of first-order autoregression; v_{-1} = lagged error term from first step of EG-model. Table based on estimates of equation (11): first step specified by equation (9) with $\theta_r = 0$, second step by equation (10). For DW, the critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about 1.156 (range of indifference from 0.747 to 1.156); see Sargan, Bhargava (1983), p. 152, Table I for T=31 and n=1. For DF, critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about -3.365, at the 10 %-level about -3.066 and at the 15 %-level about -2.864; see Phillips, Ouliaris (1990), p. 190, Table II b for n=1. Significance level of estimated coefficients (other than ϱ): '**' not significantly different from zero at the 5 %- (or higher) level; '*' significantly different from zero at the 5 %-level, but not at the 1 %-level; all other coefficients significantly different from zero at the 1 %- (or higher) level.

- albeit still not much - to be accounted for by the short-run dynamics in the second step of the EG-procedure.

(ii) As usual, the co-integration test message is mixed, with the null hypothesis (no co-integration) to be rejected at the 15 %- (or higher) level in two out of seven cases for the broad sectors and in eleven out of 31 cases for the manufacturing branches. Nevertheless, again as usual, the autoregressive parameter ρ and the coefficient α_1 as estimated in the second step of the EG-procedure remain well below 0.9, usually even below 0.8 for the vast majority of cases so that no strong conclusions should be drawn from DF- (and related) tests with apparently very little power.

(iii) By and large, the estimated structural coefficients support prior conjectures. To start with, the long-run multiplier θ_β again turns out to be close to one, although on average not quite as close as in prior estimates. Basically the same holds for the short-run coefficient β_0 of $\Delta\log(w)$ in the second step of the EG-estimate. As to the important coefficient γ_0 of the short-run disequilibrium variable $\Delta(OR^i\text{-}OR)$, the results on the whole point towards a negligible effect: for the broad sectors, γ_0 is positive in five out of seven cases, though significant at the 5 %- (or higher) level only in one, namely construction. For the 31 manufacturing branches, γ_0 has the expected sign in only eleven out of 31 cases, with none of them at a 5 %- (or higher) significance level. Of all sectors and industries, only construction and to a lesser extent banking and insurance stand out with a quite sizeable wage flexibility. In the case of construction, the familiar 10 %-drop of the industry's output ratio relative to the private sector average - remember, a spectacular cyclical event indeed - leads to a 5.5 %-cut of the average per hour earnings of employees in construction. As construction is known to be an industry with a particularly large extent of cyclical fluctuations in productivity and working conditions, this exceptional result does not come as a surprise. Note finally, that, if the employment ratio is taken as a proxy for the labour market disequilibrium, the outcome is qualitatively similar to the results obtained from estimates with the output ratio. For the broad sectors, only three out of seven cases have a positively signed coefficient γ_0, with only two being significant at the 5 %- or higher level; for the 31 manufacturing branches, the respective share is 14 out of 31, with just two of the 14 being significant at the 5 %- (or higher) level. Again, construction as well as banking and insurance stand out as sectors with a relatively high wage flexibility.

To sum up, the estimates clearly support all our prior conclusions on the high degree of interindustrial and intersectoral wage rigidity. Remember that the measure used for the wage was the broadest conceivable, at least on an hourly basis. However, one may ask the question whether elements of wage flexibility could be hidden in variations of working time so that longer-term remunerations would be more flexible than the hourly wage. In fact, this question can be tackled quite easily by reestimating the EG-model with the wage variable defined as annual (not hourly) earnings per employee.

Tables 16 and 17 present the results for these estimates in the format of the corresponding tables 14 and 15. Focusing on the main points of interest, namely the coefficients of the disequlibrium variables,[55] our prior conjectures as to the high wage rigidity are again supported, albeit with some remarkable qualifications which do point to a non-negligible cyclical role of working time as a means of earnings flexibility. For the broad sectors, the coefficient γ_0 of the disequilibrium variable $\Delta(OR^i-OR)$ is again positive in five out of seven cases, but significant at the 5 %-level in three (not just one as before), namely manufacturing, construction, and banking and insurance. For the 31 manufacturing branches, γ_0 has the expected sign in as much as 24 out of 31 cases, with 12 of them at a 5 %- (or higher) significance level. The arithmetic average of the coefficients γ_0 is in absolute terms 17.5 % for the seven sectors - 9.3 % if construction is excluded - and 8.4 % for the 31 manufacturing branches. Clearly, these are stronger estimated wage effects than those we obtained before, although they still do not indicate anything like a truely substantial wage moderation in a sector hit by a sector-specific crisis: e.g., the familiar 10 %-relative drop of a sector

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55 Note that the co-integration properties of the equations in tables 16 and 17 are more or less the same as those of the respective equations in tables 14 and 15, as a comparison of the DW-statistics and the DF-statistics shows.

Table 16 - Estimated Coefficients of the Engle/Granger-Model for Seven Sectors, Annual Earnings, West Germany 1960-1989

	long-run (1st step)							short-run (2nd step)					
	const	log w	$\bar{R}^2$	SE(%)	DW	DF	ρ	Δlog w	$\Delta(OR^i-OR)$	v_{-1}	$\bar{R}^2$	DW	$\Delta(ER^i-ER)$
Agriculture, forestry	-0.3246*	1.0088	0.9924	5.44	0.1303	-0.6306	0.9550	0.9754	0.0262**	-0.0624**	0.6194	1.2243	-0.0193**
	(0.1655)	(0.0166)						(0.0544)	(0.0587)	(0.0805)			(0.2217)
Energy, mining	-0.5056	1.0866	0.9966	3.93	0.3414	-2.1260	0.7609	1.0374	0.0613**	-0.2577*	0.7155	1.9991	0.0365**
	(0.1191)	(0.0120)						(0.0568)	(0.1290)	(0.1143)			(0.2497)
Manufacturing	-0.9792	1.1100	0.9977	3.12	0.2812	-1.6401	0.8273	1.0461	0.2982	-0.2101*	0.8677	1.1779	0.1925**
	(0.0991)	(0.0100)						(0.0366)	(0.1114)	(0.0918)			(0.1825)
Construction	0.3586	0.9606	0.9965	2.09	0.7222	-2.3197	0.6480	1.0326	0.6418	-0.3206	0.8561	1.7560	0.3029**
	(0.1076)	(0.0108)						(0.0513)	(0.1286)	(0.1176)			(0.2351)
Trade, commerce	-0.4401	1.0245	0.9985	1.87	0.5182	-3.0956	0.6605	1.0006	-0.0031**	-0.2978*	0.7191	1.3404	-0.0225**
	(0.0756)	(0.0076)						(0.0453)	(0.2761)	(0.1474)			(0.3422)
Transport, communications	0.7406	0.9402	0.9957	2.04	0.2445	-1.2698	0.8772	0.9575	-0.0462**	-0.1446**	0.7055	2.0170	-0.0202**
	(0.1153)	(0.0116)						(0.0508)	(0.2734)	(0.1082)			(0.2778)
Banking, insurance	0.1804	1.0117	0.9978	1.68	0.4521	-1.9576	0.7158	0.9449	0.2210*	-0.4025	0.7498	1.5210	0.6874
	(0.0892)	(0.0090)						(0.0418)	(0.1135)	(0.1180)			(0.1900)

Notes: $\bar{R}^2$ = adjusted R^2; SE = standard estimation error; DW = Durbin-Watson-statistic; DF = Dickey-Fuller-statistic for first-order autocorrelation of residuals; ρ = estimated coefficient of first-order autoregression; v_{-1} = lagged error term from first step of EG-model. Table based on estimates of equation (11): first step specified by equation (9) with θ_r = 0, second step by equation (10). Standard estimation errors of coefficients in parenthesis. For DW, the critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about 1.156 (range of indifference from 0.747 to 1.156); see Sargan, Bhargava (1983), p. 152, Table I for T=31 and n=1. For DF, critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about -3.365, at the 10 %-level about -3.066 and at the 15 %-level about -2.864; see Phillips, Ouliaris (1990), p. 190, Table II b for n=1. Significance level of estimated coefficients (other than ρ): '**' not significantly different from zero at the 5 %- (or higher) level; '*' significantly different from zero at the 5 %-level, but not at the 1 %-level; all other coefficients significantly different from zero at the 1 %- (or higher) level.

Table 17 - Estimated Coefficients of the Engle/Granger-Model for 31 Manufacturing Industries, Annual Earnings, West Germany 1960-1989

	long-run (1st step)							short-run (2nd step)					
	const	log w	$\bar{R}^2$	SE(%)	DW	DF	ϱ	Δ log w	$\Delta(OR^i-OR)$	v_{-1}	$\bar{R}^2$	DW	$\Delta(ER^i-ER)$
Chemicals	-0.1267	1.0417	0.9994	1.59	1.3762	-3.6722	0.3067	1.0234	0.1427	-0.6124	0.8536	1.8990	0.1845**
Refineries	0.2772*	1.0184	0.9942	5.11	0.4139	-2.5355	0.6797	1.0726	-0.0445**	-0.3135	0.5324	1.9637	-0.1382**
Plastic products	-0.1682*	1.0040	0.9983	2.70	0.5182	-2.3312	0.6951	1.0000	0.1929*	-0.3861	0.7723	1.6887	-0.1156**
Rubber	0.4620	0.9547	0.9996	1.22	1.2767	-4.1183	0.2780	0.9749	0.0544**	-0.7730	0.8952	1.6450	0.0498**
Stone and minerals	0.3274	0.9769	0.9989	2.09	0.7656	-2.5871	0.5879	0.9924	0.2227	-0.3389	0.8610	1.6428	0.3961
Ceramics	0.2770	0.9543	0.9983	2.60	0.3385	-1.9372	0.7823	0.9922	-0.0145**	-0.2637*	0.8442	1.8960	-0.2622*
Glass	-0.0892**	1.0062	0.9987	2.35	0.4843	-1.7726	0.7636	1.0030	0.1142**	-0.2825*	0.7802	1.7731	0.2548*
Iron and steel	0.7581	0.9425	0.9967	3.57	1.8398	-5.3078	-0.0202	0.9672	0.0728**	-0.9964	0.6453	2.0360	0.5854*
Non-ferrous metals	0.0455**	1.0051	0.9990	2.13	0.6477	-2.3736	0.6562	1.0008	-0.0456**	-0.3810*	0.8403	1.8679	-0.1946*
Foundries	0.5077	0.9552	0.9993	1.67	1.2339	-3.7032	0.3161	0.9713	0.1359	-0.5556	0.8911	2.2201	0.1520*
Rolling mills	0.7335	0.9191	0.9993	1.64	0.7332	-2.3735	0.6352	0.9511	0.2274	-0.2513**	0.9322	2.2291	0.1612**
Metal manufacturing	0.9885	0.9015	0.9960	3.76	0.2895	-1.3797	0.8552	0.9616	0.0001**	-0.2291*	0.7976	1.1473	0.0235**
Mechanical engineering	0.1693	0.9915	0.9993	1.73	0.6747	-2.0542	0.6727	1.0078	0.1113**	-0.3601*	0.8828	1.9858	0.0079**
Business machines	-2.0444	1.2370	0.9982	3.42	1.1647	-2.9397	0.3920	1.2816	0.0592**	-0.6708	0.6538	2.0213	0.0089**
Motor vehicles	-0.3103	1.0417	0.9992	1.96	0.9561	-3.1811	0.4591	1.0132	0.1686	-0.5046	0.8663	2.0379	0.1434**
Shipbuilding	0.2732	0.9824	0.9943	4.89	0.4812	-2.2158	0.7166	1.0699	0.1137	-0.3117	0.6493	1.7423	0.2729*
Air- and spacecraft	-1.1948	1.1395	0.9982	3.20	0.6245	-2.1617	0.6998	1.1436	-0.0019**	-0.3368	0.6421	1.6192	0.0647**
Electrical engineering	-0.5128	1.0537	0.9996	1.34	1.7655	-4.8389	0.0565	1.0203	-0.1339**	-0.9570	0.8903	2.1074	-0.1886*
Optical instruments	0.1410	0.9712	0.9995	1.47	0.6624	-2.2438	0.6773	0.9535	0.0523**	-0.3118*	0.8700	2.0834	-0.0729**
Iron, tin and metal goods	-0.1977	1.0142	0.9994	1.64	0.6081	-2.2057	0.6891	1.0061	0.1538*	-0.4016	0.8842	1.5193	0.1538*
Musical instruments and toys	-0.6684	1.0407	0.9978	3.20	0.6485	-2.0893	0.6891	0.9700	0.1766*	-0.3077*	0.6557	2.2023	-0.2294**
Wood working	-0.2304	1.0164	0.9987	2.36	0.5627	-2.5842	0.6442	1.0274	0.0914**	-0.4168	0.7594	1.8575	-0.0716**
Wood processing	1.1281	0.8672	0.9969	3.17	0.3584	-1.8831	0.7793	0.9046	0.1480**	-0.2060*	0.8013	2.1580	0.1504**
Cellulose, pulp and paper	-0.0139**	1.0151	0.9983	2.74	0.8474	-3.1923	0.5285	1.0200	0.1486*	-0.3949*	0.7505	2.1639	-0.1290**
Pulp and paper processing	-0.4435	1.0280	0.9998	0.98	0.9066	-3.1469	0.4707	1.0325	0.0336**	-0.5808	0.9391	1.8017	-0.0164**
Printing	0.7388	0.9216	0.9963	3.69	0.1592	-0.8499	0.9290	0.9263	-0.0806**	-0.0576**	0.8242	1.6801	-0.1846**
Leather	0.5780	0.9089	0.9939	4.64	0.2283	-0.9666	0.9052	0.8621	0.2132*	-0.0879**	0.5503	2.0227	0.0131**
Textiles	-0.2093	0.9955	0.9997	1.12	0.7757	-2.3606	0.6182	0.9809	-0.0760**	-0.3841*	0.9067	1.6627	-0.2285*
Clothing	0.0886	0.9432	0.9996	1.25	0.9800	-2.8821	0.5107	0.9309	0.1825	-0.3991*	0.9135	2.0974	0.1522**
Food and beverages	1.6816	0.8143	0.9990	1.73	0.5536	-2.3544	0.6531	0.8232	0.1563	-0.3005*	0.8699	1.8840	0.2022*
Tobacco	-2.8108	1.2780	0.9985	3.28	1.1247	-3.1914	0.4192	1.2527	0.0408**	-0.6027*	0.5338	1.6169	-0.6124*

Notes: $\bar{R}^2$ = adjusted R^2; SE = standard estimation error; DW = Durbin-Watson-statistic; DF = Dickey-Fuller-statistic for first-order autocorrelation of residuals; ϱ = estimated coefficient of first-order autoregression; v_{-1} = lagged error term from first step of EG-model. Table based on estimates of equation (11): first step specified by equation (9) with $\vartheta_r = 0$, second step by equation (10). For DW, the critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about 1.156 (range of indifference from 0.747 to 1.156); see Sargan, Bhargava (1983), p. 152, Table I for T=31 and n=1. For DF, critical value for rejecting the null hypothesis of a unit root in the residuals is at the 5 %-level about -3.365, at the 10 %-level about -3.066 and at the 15 %-level about -2.864; see Phillips, Ouliaris (1990), p. 190, Table II b for n=1. Significance level of estimated coefficients (other than ϱ): '**' not significantly different from zero at the 5 %- (or higher) level; '*' significantly different from zero at the 5 %-level, but not at the 1 %-level; all other coefficients significantly different from zero at the 1 %- (or higher) level.

output ratio would lead on average to a 0.84 %-relative wage cut in the manufacturing branches and a 1.74 % wage cut in the broad sectors (0.93 % if construction is excluded). Again, only construction stands out with a sizeable wage flexibility, implying a 6.4 %-relative wage cut in case of the above crisis event.

If the employment ratio is chosen as the disequilibrium proxy, the results are similar though much less conclusive than with the output ratio: for the broad sectors, four out of seven estimates of γ_0 are positive, with just one of them being significant at a 5 %- (or higher) level; for the 31 manufacturing branches, the respective share is 19 out of 31 with seven of them being significant at the 5 %- (or higher) level. Somewhat surprisingly, it is now banking and insurance - not construction - which stands out with the greatest extent of wage flexibility.

6. Concluding Remarks

Our analysis has shown that structural wage rigidity is a pervasive feature of labour markets in Germany: neither between regions nor between sectors and industries do wages react to any significant degree to respective labour market disequilibria, be they temporary or permanent. Minor exceptions - for regions: the isolated city-state of West Berlin before German unification, for sectors: the construction industry - are worth noting, but do clearly not disprove the rule. The high structural wage rigidity shows up both in contractual minimum and in effective wages, albeit slightly more so in the former than in the latter, which points to at least a modest role for wage drift as a means to achieve some post-contractual flexibility of the wage structure. If longer-term earnings such as annual remunerations are considered, variations in working time seem to add another element of flexibility, though again, it is a rather modest one.

As always in econometrics, our results hinge upon the data used and the methodology applied; they are thus subject to the many qualifications discussed in the paper. Nevertheless, there are at least two good reasons to believe that the main message of the results is fairly robust.

First, the empirical picture has been remarkably uniform and unambiguous, with virtually all estimates pointing towards high structural wage rigidity independent of the data source and the model specification. Second, the long-run equilibrium relationship and thus the long-term co-movements of wages between different structural units turned out to be very tight all throughout, in particular between regions. This means that, even if the short- and medium-run disequilbrium dynamics has been repeatedly misspecified, one would still have to conclude that the structural wage effects of these disequilibria are likely to be rather small - simply because not much 'autonomous' structural wage dynamics is to be explained anyway. Hence the least controversial message of our estimates by itself indicates a rather limited structural wage flexibility.

What are we to make of our results for economic policy in post-unification Germany? The main conclusion looks quite obvious: for whatever reason, German industrial unions and employers' associations, who are the main agents at the wage bargaining table, regard structural disequilibria as being beyond their competence and responsibility; in turn, market forces - working via wage drift - are apparently insufficient to loosen the straightjacket of a rigid wage structure to any substantial degree. If this has been so in the last four decades, why should one expect it to be different in an economically unified Germany which has basically taken over the whole institutional framework of collective bargaining from prior West Germany? In this sense, German unification may simply have redefined the geographical entity over which wage rigidity is to prevail. Hence it should not be all that surprising to see collectively agreed minimum wages in East Germany - and, in their tow, effective wages as well - quickly move up to the level of the West as there are in fact set to do in most industries up to the mid-1990s. The huge industrial transformation, in which the economy of the East finds itself, is thus implicitly interpreted by unions and employers' associations as nothing but another regional crisis with costs of adjustment that have to be largely borne by the state through a drastic subsidization of capital investment, labour mobility and regional unemployment. Such-like crises have happened many times in the West, although, of course, to a much less dramatic extent. Thus interpreted, the efforts to equalize the wage between West and East are simply the immediate consequence of a traditional philosophy, with the only (gradual) difference being that the regional disequilibrium is of a vastly greater dimension than at any time before.

A change of this philosophy could only be induced if the government refrained from socializing the costs of structural crises and thus forced unions and employers' associations to take over full responsibility for mitigating any industrial or regional downturn. In fact, the sheer size of the crisis in East Germany and the concommittant huge burden for public finances would have given a unique historical opportunity to put moral pressure on collective bargaining to allow for a substantial interregional wage differentiation and thus to initiate a more self-sustained and less government-sponsored industrial transformation. For whatever reason, this chance was missed.

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